Deforestation, agriculture expansion and agricultural productivity in tropical Latin America: is there a Jevons paradox?

M.G. Ceddia[1ł], S. Sedlacek[ł], N.O. Bardsley[ǂ] and S. Gomez-y-Paloma[§]

ł MODUL University Vienna, Department of Public Governance, Austria

ǂ University of Reading, School of Agricultural Policy and Development, UK

§ European Commission Joint Research Centre, IPTS, Spain

Abstract

The process of global deforestation calls for urgent attention, particularly in South America where deforestation rates have failed to decline over the past 20 years. The main direct cause of deforestation is land conversion to agriculture, while the underlying drivers include economic growth, population growth and other technological and institutional factors. As world population is expected to reach 9 billion by 2050, the scientific and policy communities are placing significant emphasis on technological innovation and agricultural intensification in order to address the food security issue without further impacting on existing forests. In this article we combine data from the FAO and the World Bank for 7 Southern American countries over the period 1970-2006 estimate an econometric model which accounts for various determinants of agricultural land expansion and use elasticities to quantify the effect of the different independent variables. In particular we investigate whether, after accounting for

[1] Corresponding author. Email: graziano.ceddia@modul.ac.at

various socio-economic factors and corruption control, an increase in agricultural productivity behaves unexpectedly by strengthening the incentives to expand agricultural area, therefore leading to a "Jevons paradox". Our results show that when corruption control is strong a Jevons paradox begins to emerge for high levels of agricultural productivity. We also find that agricultural expansion is positively related to the level of service on external debt, agricultural exports and population growth. Finally, the Environmental Kuznets Curve hypothesis for agricultural area, suggesting that GDP growth initially promotes agricultural expansion but ultimately leads to a reduction in agricultural land, is not supported by the data. Instead the effect of per-capita GDP on agricultural land is highly non-linear but ultimately leads to an increase in agricultural area.

Introduction

The FAO has documented how over 2000-2010 more than 50 million ha of forest have been lost globally, with an annual loss rate of -0.13%. At the global level the process of deforestation has slowed down, from 1990-2000 when forests were being lost at a rate of -0.20% per annum, but this overall trend hides important regional differences. Central and South America account for over 20% of the world's forest extension. The annual rate of deforestation over the decades 1990-2000 and 2000-2010 has been -1.56% and -1.19% for Central America, while it has remained constant at -0.45% for Southern American (FAO, 2011). These rates, although seem to be slowing down (at least in Central America), remain substantially higher than those observed at the global level. In order to properly understand the process of deforestation in Southern America (and elsewhere), it is important to distinguish two different causative mechanisms: proximate and underlying causes (Millennium Ecosystem Assessment, 2005; Geist & Lambin, 2002). Proximate causes refer to activities or actions at the local level. Evidence suggests that the main proximate cause of deforestation in tropical regions is agricultural expansion, followed by infrastructure expansion and commercial wood extraction (Grau & Aide, 2008; Bawa & Dayanadan, 1997). Underlying causes reflect fundamental socio-economic, technological and political/institutional factors which drive the local proximate causes. With respect to the socio-economic factors, some of the determinants of agricultural expansion include high cash-crop prices, low production costs and the need to generate foreign exchange earning in order to service external debt (Geist & Lambin, 2002; Shandra et al., 2008; Bawa & Dayanadan, 1997). The effect of an increase in per-capita GDP has been investigated at length, particularly to test the hypothesis that an initial increase in GDP will stimulate agricultural expansion only up to a certain point beyond which a further increase in income will lead to a reduction in agricultural area (Barbier and Burgess, 2001; Barbier, 2004a,b). Such a hypothesis, known as Environmental Kuznets Curve in the environmental economic literature, is very attractive as it suggests that a country can "grow" itself out of environmental degradation. The effect of population increase on the expansion of agricultural area has also been ascertained (Bawa & Dayanadan, 1997; Barbier and Burgess, 2001; Barbier, 2004a,b), plausibly because of its effect on food demand (Von Braun, 2007; Deutsche Bank, 2009). With respect to the role of political-institutional factors, it has been shown how better corruption control and increased political stability are associated with agricultural expansion (Barbier and Burgess, 2001; Barbier, 2004a,b). Lower corruption improves bureaucratic efficiency and facilitates land credit and policy which stimulate the conversion of forests to agriculture (Lopez, 1998).

As the world population is expected to reach 9 billion by 2050, food security has become a priority issue on the international political agenda. Given the prominent role played by agricultural expansion on tropical deforestation an intense scientific debate has emerged on how to feed an increasing population without further damaging existing forests. In this context the debate over the role of technological innovation, relieving some of the pressures associated with agricultural expansion, has been mainly framed with regard to the benefits of sustainable intensification (Royal Society, 2009) and the phenomenon of “land-sparing” (Ewers et al., 2009). These are certainly powerful ideas, and part of their attractiveness is related to their simplicity: if one can produce more per unit of land then one certainly needs less land (Meyfroidt & Lambin, 2007). From a theoretical point of view, it has been noted that the land-sparing effect of agricultural intensification is likely to emerge if the demand for agricultural products is inelastic, as the increased production will significantly lower prices, reduce revenues and weaken the incentives for further land conversion (Bashaasha et al., 2001). However, it has also been noted how intensification and yield increase may produce a “Jevons paradox” and strengthen the incentives to further expand agricultural area (Lambin & Meyfroidt, 2011; Rudel et al., 2009; Coxhead et al., 2001; Kaimowitz & Smith, 2001).

The remainder of the paper is structured as follows. After the present introduction, we discuss the data and the methods employed in the analysis. Subsequently we present and discuss our results. Finally, we draw the conclusions.

Data and methods

Drawing on FAO and World Bank data, here we conduct an econometric analysis to assess whether agricultural intensification has been promoting or deterring agricultural expansion in 7 countries in Latin America (Bolivia, Brazil, Colombia, Ecuador, Paraguay, Peru and Venezuela) over the period 1970-2006. Agricultural intensification is quantified as the value of agricultural output (at constant prices) per ha of agricultural land (Figure 1.a), while the dependent variable is the agricultural area in the i-th country at time t (Figure 1.b).

[Figure 1 here]

We also account for various socio-economic factors (per-capita GDP, agricultural exports, agricultural value added, population and service on external debt) and corruption control and form two nested econometric models (one including the corruption control index). The models are estimated by running a robust panel data regression with both one-way and two-way random and fixed effects.

A panel data set consists of a repeated cross-section, involving individual units i=1….N, over a certain period of time t=1…T. One of the main advantages of using panel data is the ability to control for a number of unobserved factors, including time-invariant factors (i.e., that vary across individuals but not over time) and individual-invariant factors (i.e., that vary across time but are the same for all individuals). A detailed description of the various methods to estimate panel data can be found in Wooldridge (2002). Here we only provide a short description of the methods employed in the article.

Given a general econometric model of the form

$$y_{it} = \alpha + \beta_1 x_{it1} + \cdots + \beta_k x_{itk} + v_{it} \quad \text{(1.a)}$$

$$v_{it} \sim IID(0, \sigma^2) \quad \text{(1.b)}$$

The existence of time-invariant individual-specific unobserved effects implies that (1.a) can be rewritten as

$$y_{it} = \alpha_i + \beta_1 x_{it1} + \cdots + \beta_k x_{itk} + v_{it} \quad \text{(2)}$$

Where the constant α_i varies across individuals but is constant over time. Expression (2) can be estimated either through the one-way Fixed Effect model (FE) or through the one-way Random Effect model (RE). The FE assumes that α_i are just fixed constant. The model is normally estimated by running an ordinary least square (OLS) regression of $(y_{it} - \bar{y}_{it})$ on $(x_{itk} - \bar{x}_{itk})$, where $\bar{y}_{it}$ and $\bar{x}_{itk}$ represent the individual-specific mean of y_{it} and x_{itk} respectively, and subsequently recovering the α_i intercepts. The RE specification assumes that (2) can be expressed as $y_{it} = \mu + \beta_1 x_{it1} + \cdots + \beta_k x_{itk} + \alpha_i + v_{it}$ with $\alpha_i \sim \left(0, \sigma^2_{\alpha_i}\right)$, where the individual-specific effects α_i are stochastic. The RE model is estimated through a Generalized Least Square (GLS) procedure. This involves running an OLS regression of $(y_{it} - \theta\bar{y}_{it})$ on $(x_{itk} - \theta\bar{x}_{itk})$, which implies transforming the original data by removing a fraction θ (this parameter is also estimated) of the individual-specific means (see Wooldridge 2002 for details).

When individual-invariant (i.e., time-specific) effects are also included (2) can be rewritten as

$$y_{it} = \alpha_i + \lambda_t + \beta_1 x_{it1} + \cdots + \beta_k x_{itk} + v_{it} \quad (3)$$

Where λ_t indicates the time-specific effect. Expression (3) can also be estimated by assuming that the individual-specific components α_i and the individual-invariant components λ_t are constants (obtaining the two-way FE model) or stochastic (obtaining the two-way RE model).

The choice between ordinary least square (OLS) of a pooled cross-section against one-way and two-way FE and RE models is performed on the basis of the (Breusch-Pagan) Lagrange Multiplier test (a large value of this statistic argues in favor of the FE or RE models against OLS) and Hausman test (a large value of this statistic argues in favor of the FE over the RE model).

In order to address the question of whether in Southern America agricultural intensification has allowed for land-sparing or on the contrary has further stimulated agricultural expansion, therefore revealing a Jevons paradox, we start from the following general econometric model

$$\mathrm{y_{it}} = \alpha_i + \lambda_t + \boldsymbol{\beta}\mathbf{x_{it}} + \boldsymbol{\theta}\mathbf{z_{it}} + v_{it} \quad (4)$$

The dependent variable in expression (4) represents agricultural land in the i-th country at time t (AL_{it}), while the independent variables reflect a vector of technological/intensification indicators ($\mathbf{x_{it}}$) and socio-economic variables ($\mathbf{z_{it}}$). The next step is to incorporate corruption control into (4). First, notice that given the limitation of public available data we consider an index of corruption control ($\mathbf{q_i}$) which changes across countries but not over time. Second, we are particularly interested in analyzing the interaction of corruption control with one of the intensification factors ($\mathbf{x_{it}}$). We then construct a new variable $\mathbf{w_{it}}=\mathbf{q_i}\times\mathbf{x_{it}}$ to obtain the following model

$$\mathrm{y_{it}} = \alpha_i + \lambda_t + \boldsymbol{\gamma}\mathbf{w_{it}} + \boldsymbol{\beta}\mathbf{x_{it}} + \boldsymbol{\theta}\mathbf{z_{it}} + v_{it} \quad (5)$$

In expressions (4) and (5) α_i and λ_t represent the individual-specific and time-specific effects while $\boldsymbol{\gamma}$, $\boldsymbol{\beta}$ and $\boldsymbol{\theta}$ are vectors of parameters associated with the corruption, technological/intensification and socio-economic variables respectively, while v_{it} is the independently distributed error term with zero mean and finite variance.

Drawing on FAO data on agricultural land for 7 Latin American countries (Bolivia, Brazil, Colombia, Ecuador, Paraguay, Peru and Venezuela) over the period 1970-2006, the dependent variable in expressions (4) and (5) is formed as the logarithm of agricultural land in the i-th country at time t. In order to construct a vector of technological/intensification variables we use FAO data on the value of agricultural output (at constant prices) and divide it by the corresponding agricultural land indicator to obtain a measure of value of agricultural output (at constant prices) per unit of land ($AOHA_{it}$). The elements of the vector $\mathbf{x_{it}}$ used for the empirical analysis include $log(AOHA_{it})$ and $log^2(AOHA_{it})$. To form a vector of socio-economic factors, we rely on data from the World Bank concerning population (POP_{it}), per-capita income at constant 2000 US$ ($GDPC_{it}$), agricultural value added as a share of the total GDP (AVA_{it}), service on external debt as a share of the GDP at current prices ($PEDS_{it}$) and FAO data on export quantity index (EX_{it}). The elements of the vector $\mathbf{z_{it}}$ used for the empirical analysis include $log(POP_{it})$, $log^2(POP_{it})$, $log(GDP_{it})$, $log^2(GDP_{it})$, $log^3(GDP_{it})$, $log(EX_{it})$, $log^2(EX_{it})$, $log(AVA_{it})$, $log(PEDS_{it})$, $log^2(PEDS_{it})$, $log(PEDS_{it}) \times log(AVA_{it})$.

In order to account for the level of corruption, we use the World Bank corruption control index. The index, which varies between -2.5 (no corruption control) and +2.5 (high corruption control) is available for all the countries included in our analysis but only starting from 1996. As we need a measure of corruption control which could be applied to the entire horizon covered by our dataset (1976-2006) we compute the country-specific (and time-invariant) average of the corruption control index over the period 1996-2006 ($CORC_i$) and use it to form the vector $\mathbf{q_i}$. Finally, to test whether the level of corruption interacts with agricultural intensification we form the vector w_{it} as $CORC_i \times AOHA_{it}$.

Our sample initially contains 37×7=259 observations, but after dropping missing values we are left with 217 observations. The descriptive statistics are presented in Table 1.

[Table 1 here]

From expressions (4) and (5) we obtain the following pair of nested models (subsequently referred to as model 1.a and 1.b respectively)

$$log(AL_{it}) = \alpha_i + \lambda_t + \beta_1 log(AOHA_{it}) + \beta_2 log^2(AOHA_{it}) + \theta_1 log(POP_{it}) + \theta_2 log^2(POP_{it}) + \theta_3 log(GDPC_{it}) + \theta_4 log^2(GDPC_{it}) + \theta_5 log^3(GDPC_{it}) + \theta_6 log(EX_{it}) + \theta_7 log^2(EX_{it}) + \theta_8 log(PEDS_{it}) + \theta_9 log^2(PEDS) + \theta_{10} log(AVA_{it}) + \theta_{11}[log(PEDS_{it}) \times log(AVA_{it})] + v_{it} \quad (6.a)$$

$$log(AL_{it}) = \alpha_i + \lambda_t + \gamma_1(CORC_{it} \times AOHA_{it})\beta_1 log(AOHA_{it}) + \beta_2 log^2(AOHA_{it}) + \theta_1 log(POP_{it}) + \theta_2 log^2(POP_{it}) + \theta_3 log(GDPC_{it}) + \theta_4 log^2(GDPC_{it}) + \theta_5 log^3(GDPC_{it}) + \theta_6 log(EX_{it}) + \theta_7 log^2(EX_{it}) + \theta_8 log(PEDS_{it}) + \theta_9 log^2(PEDS) + \theta_{10} log(AVA_{it}) + \theta_{11}[log(PEDS_{it}) \times log(AVA_{it})] + v_{it} \quad (6.b)$$

Results

Given the functional specification adopted, most of the parameters reflect elasticities. In fact, this measure is the most appropriate when the independent variables are expressed in different units (Cariboni et al., 2007). Given a function $f(x_1 \ldots . x_n)$ the elasticity with respect to the k-th variable is defined as $\varepsilon_{x_k} = \frac{\partial f(\cdot)/f(\cdot)}{\partial x_k/x_k} \equiv \frac{\partial log[f(\cdot)]}{\partial log(x_k)}$ and indicates the % change in the dependent variable for a 1% change in the k-th independent variable. The presence of logarithm products among the explanatory variables in expressions (6.a) and (6.b) implies that the elasticities are not constant. The use of this kind of logarithmic specification is common in the environmental-economic literature, particularly with reference to the Environmental Kuznets Curve hypothesis (e.g., Stern 2010; Stern and Common, 2001).

The estimation of the models is performed by running a robust panel data regression with both one-way and two-way random and fixed effects. The main results are illustrated in Table 2.

[Table 2 here]

The performance of the nested models (6.a,b) is assessed on the basis of the likelihood ratio test $LR = 2\big(LogLik(modelB) - LogLik(modelA)\big)$ with $LR \sim \chi^2(m)$, where m indicates the number of additional variables in model B, and the Akaike Information Criterion (AIC). We obtain LR=22.307 and m=1 ($p<0.001$). This implies that model (6.b) performs significantly better than (6.a). However, when the AIC is used, we cannot easily discard the results associated with (6.a). As reported in Table 2, the lowest AIC is associated with expression (6.b), which corresponds to model (1.b) in the table. Burnham and Anderson (2002, p. 70) have noted that when differences between the AIC of various models and the minimum AIC (i.e., the one associated with model 1.a in our case) is less than 2, then none of the models estimated can be deemed to be superior. In our case, given the very small differences between the AIC, we believe that all models should in general be considered valid.

Consider the model without the corruption control index first (models 1.a in Table 2). The model explains a significant portion of the overall variation. The elasticity of agricultural land with respect to agricultural intensification is $\varepsilon_{AOHA} = \beta_1 + 2\beta_2 log(AOHA)$. In model (1.a) β_1=-0.401 ($p<0.001$) while β_2 is not significantly different from zero, implying that ε_{AOHA} is constant and negative and that intensification always promotes land-sparing (figure 2). As for the effects of the various socio-economic variables we find that in general the coefficients in model (1.a) have the expected sign. The elasticity with respect to population is $\varepsilon_{POP} = \theta_1 + 2\theta_2 log(POP)$ with θ_1=0.756 ($p<0.1$) and θ_2 not significantly different from zero. This suggests that an increase in population is likely to promote agricultural expansion, a result which is in line with previous findings (Lambin, Geist and Lepers, 2003; Bawa and Dayanadan, 1997). The elasticity of agricultural land with respect to per-capita GDP is $\varepsilon_{GDPC} = \theta_3 + 2\theta_4 log(GDPC) + 3\theta_5 log^2(GDPC)$. The estimated value of the relevant parameters are θ_3=59.519 ($p<0.001$), θ_4=-7.882 ($p<0.001$) and θ_5=0.347 ($p<0.001$) in model (1.b). This suggests that ε_{GDPC} is initially positive, until a first "turning point" (τ_1) is reached, then becomes negative until a second "turning point" (τ_2) is reached and thereafter remains positive (figure 3). An increase in per-capita income is initially associated with an expansion of the agricultural area, subsequently a further income increase leads to a reduction in the agricultural area, but after crossing the second turning point an increase in per-capita income ultimately promotes a further expansion of the agricultural area. Previous evidence suggested the existence of only one turning point, supporting the Environmental Kuznets Curve hypothesis: initially the increase of per-capita income is likely to promote environmental degradation and agricultural expansion but as a country becomes more affluent it will "grow-out" of environmental degradation and the agricultural area will shrink (Barbier and Burgess, 2001; Barbier 2004a,b). Our results suggest that the relationship between per-capita income and agricultural expansion could be more complex, something which has recently been pointed out with respect to other environmental impacts (Strong et al., 2011). On the basis of the estimated parameter values, for model (1.a) the turning points are τ_1=1,312 US$ per person and τ_2=2,875 US$ per person (see the Appendix for details on their computation). These points are all internal to the sample. All the other coefficients (except the constant) are not significantly different from zero.

[Figure 2 here]

Consider now model (1.b), where an index of corruption control interacting with the agricultural intensification variable is introduced among the explanatory variables. This model also explains a significant portion of the overall variations. The inclusion of the institutional variables reflecting corruption control and its interaction with intensification significantly affects the results. The elasticity of the agricultural area with respect to intensification is now $\varepsilon_{AOHA} = \gamma_1(CORC \times AOHA) + \beta_1 + 2\beta_2 log(AOHA)$. In model (1.b) we obtain γ_1=0.001 ($p<0.001$), β_1=-0.709 ($p<0.001$) and β_2=0.073 ($p<0.001$). To fully appreciate these results we plot ε_{AOHA} as estimated from (1.b) for three different values of the corruption control index corresponding respectively to the sample mean, the sample minimum and the sample maximum (figure 2). When corruption control is low or modearate (i.e., CORC is evaluated at the sample minimum or sample mean), ε_{AOHA} remains negative suggesting that an increase in agricultural productivity produces a land-sparring effect. However, when the corruption control index is high (CORC is evaluated at its sample maximum) ε_{AOHA} is increasing and for a large enough increase in productivity it becomes positive, suggesting the emergence of a Jevons paradox. As it has been noted before, corruption control may increase bureaucratic efficiency and therefore facilitate agricultural expansion (Lopez, 1998). This result also implies that the land-sparing effect of agricultural intensification should be carefully considered, on a case-by-case basis as the supporting evidence remains mixed. The other socio-economic factors behave as expected. The effect of population growth on agricultural expansion is clearly positive in model (1.b). The elasticity of the agricultural area with respect to per-capita income ε_{GDPC} follows a pattern similar to the one associated with models (1.a), signaling the existence of two "turning points" (Figure 3). Initially ε_{GDPC} is positive and an increase in income leads to an expansion of the agricultural area, until τ_1 is reached; after this point ε_{GDPC} is negative and an increase in income generates a reduction of agricultural area until τ_2 is reached; beyond this point ε_{GDPC} is again positive and income increase ultimately promotes the expansion of agriculture. On the basis of the estimated parameters for model (1.b) the turning points are τ_1=1,121 US$ per-capita and τ_2=3,516 US$ per-capita.

[Figure 3 here]

Consider next the effects of the service on external debt. For model (1.b) we have $\varepsilon_{PEDS} = \theta_8 + 2\theta_9 log(PEDS) + \theta_{11} log(AVA)$, with θ_8=-0.118 ($p<0.1$), θ_9 not significantly different from zero and θ_{11}=0.054 ($p<0.005$). An increase in the service on external debt promotes agricultural expansion, only in those countries with a “sufficiently large” agricultural sector. On the basis of the estimated coefficients we find that the service on external debt begins to have a positive effect on the expansion of agricultural area when at least 8.9% of the GDP derives from the agricultural sector (Figure 4).This result, indicating a sort of path dependence, is to some extent corroborated by existing evidence that suggests how countries in which the agricultural sector plays a minor role are less prone to further expand agriculture as a under the pressure of external debt (Shandra et al., 2008). Finally, the effect of agricultural exports is not statistically significant.

Conclusions

Our econometric analysis provides an assessment of the effect of agricultural intensification, alongside other relevant socio-economic determinants, on agricultural expansion in Southern America over the period 1970-2006. Our results show that when corruption control is not accounted for the land-sparing effect of intensification is clearly detected. However, when corruption is accounted for, a Jevons paradox emerges when corruption control is more intense. Drawing on this result we would like to recommend caution when identifying agricultural intensification as the obvious mean by which to reduce agriculture expansion in Southern America, since it may interact in unexpected ways with other policy drivers, including those aiming at lowering corruption levels. In particular, while we recognize that promoting agricultural intensification may be necessary to address food security, preventing the further expansion of agricultural areas will also require a specific policy intervention, like the introduction of some capping mechanism. The effect of population growth and service on external debt also appears to be consistent with both expectations and previous results. As for the role of per-capita GDP, our results reject the existence of an Environmental Kuznets Curve hypothesis and indicate that an increase in affluence is likely to ultimately lead to a further agricultural expansion. In this respect, we also recommend that any policy that aims at promoting economic growth as a way to limit the expansion of agriculture should be carefully assessed.

References

Barbier EB (2004a). Explaining agricultural land expansion and deforestation in Developing Countries. American Journal of Agricultural Economics 86(5): 1347-1353.

Barbier EB (2004b). Agricultural expansion, resource booms and growth in Latin America: implications for long-run economic development. World Development 32(1): 137-157.

Barbier EB and Burgess JC (2001). The economics of tropical deforestation. Journal of Economic Surveys 15(3): 413-433.

Bashaasha B, Kraybill DS and Southgate DD (2011). Land use impacts of agricultural intensification and fuelwood taxation in Uganda. Land Economics 77(2): 241-249.

Bawa KS and Dayanadan S (1997). Socioeconomic factors and tropical deforestation. Nature 386: 562-563.

Burnham KP and Anderson DR (2002). Model selection and multimodel inference: a practical information-theoretic approach, 2^{nd} Edition, Springer.Cariboni et al. (2007).

Coxhead I, Shively G and Shuai X (2001). Agricultural development policies and land expansion in southern Philippines watershed. In Angelsen A and Kaimowitz D (Ed.), Agricultural technologies and tropical deforestation, p. 347-366, CABI, CIFOR.

Deutsche Bank (2009). The Global Food Equation. Deutsche Bank Research, available at www.dbresearch.com.

Ewers RM, Scharlemann JPW, Balmford A and Green RE (2009). Do increase in agricultural yield spare land for nature? Global Change Biology 15(7): 1716-1726

Food and Agriculture Organization (2011). State of the world's forests. FAO, Rome.

Giest HJ and Lambin EF (2002). Proximate causes and underlying driving forces of tropical deforestation. Bioscience 52(2): 143-150.

Grau HL and Aide M (2008). Globalization and land-use transition in Latin America. Ecology and Society 13(2).

Kaimowitz D and Smith J (2001). Soybean technology and the loss of natural technology in Brazil and Bolivia. In Angelsen A and Kaimowitz D (Ed.), Agricultural technologies and tropical deforestation, p. 195-212, CABI, CIFOR.

Lambin EF, Geist HJ and Lepers E (2003). Dynamics of land-use and land-cover change in tropical regions. Annual Review of Environment and Resources 28: 205-241.

Lopez R (1998). Where development can or cannot go: the role of poverty-environment linkages. In Pleskovic B and Stiglitz JE (Ed.), Annual Bank Conference on Development Economics 1997, p. 285-306, The World Bank, Washington DC.

Meyfroidt P and Lambin EF (2007). The causes of reforestation in Vietnam. Land Use Policy 25: 189-207.

Millennium Ecosystem Assessment (2005). Ecosystems and human wellbeing: synthesis. Island Press, Washington, DC.

Royal Society (2009). Reaping the benefits. Science and the sustainable intensification of global agriculture.

Rudel TK, Schneider L, Uriarte M, et al. (2009). Agricultural intensification and changes in cultivated areas, 1970-2005. Proceedings of the National Academy of Sciences 106(49): 20675-20680.

Shandra JM, Shor E, Maynard G and London B (2008). Debt, structural adjustment and deforestation: a cross-national study. Journal of World-Systems Research, 14(1): 1-21.

Stern DI (2010). Between estimates of the emissions-income elasticity. Ecological Economics 69: 2173-2182.

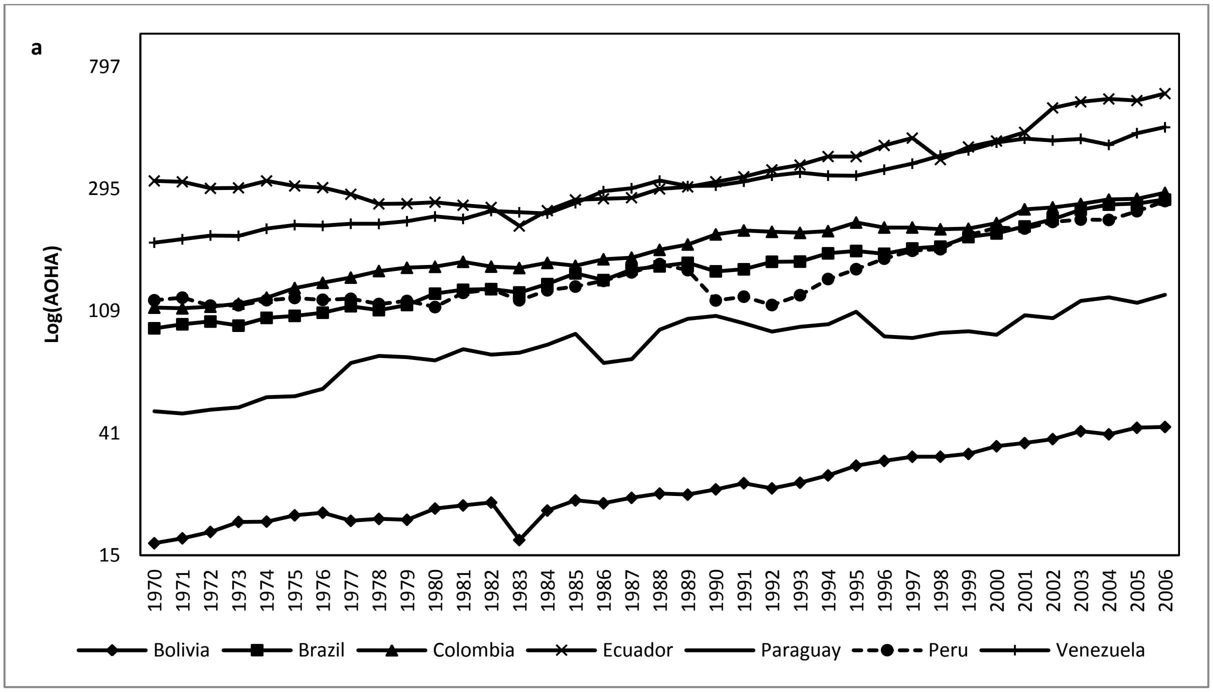
Stern DI and Common M (2001). Is there an environmental Kuznets curve for sulfur? Journal of Environmental Economics and Management 41: 162-178.

Strong A, Tschirhart J and Finnoff D (2011). Is economic growth for the birds? Ecological Economics 70: 1375-1380.

Von Braun J (2007). The world food situation. New driving forces and required actions. IFPRI Food Policy Report.

Wooldridge JM (2002). Econometric analysis of cross section and panel data. The MIT Press, Cambridge (Massachusetts), London (England).

Figures and Tables



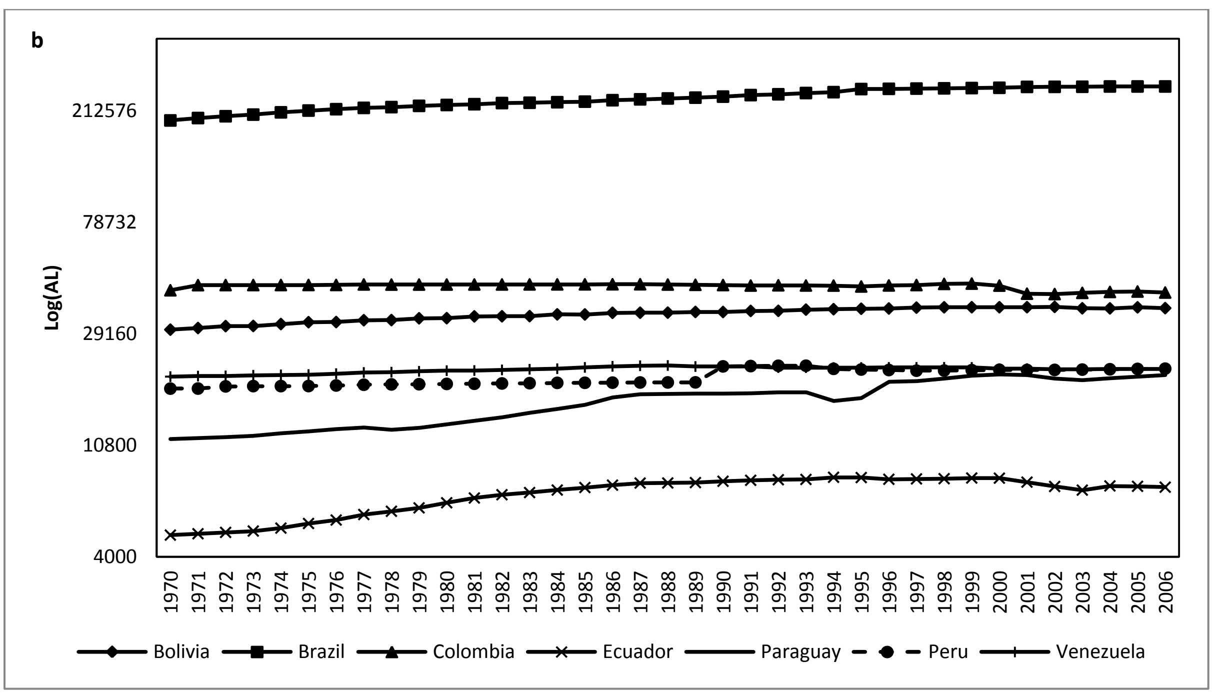


Figure 1: (a) the evolution of agricultural intensification, as measured by the value of agricultural output (at constant prices) per ha (AOHA). **(b)** The evolution of agricultural land (AL).

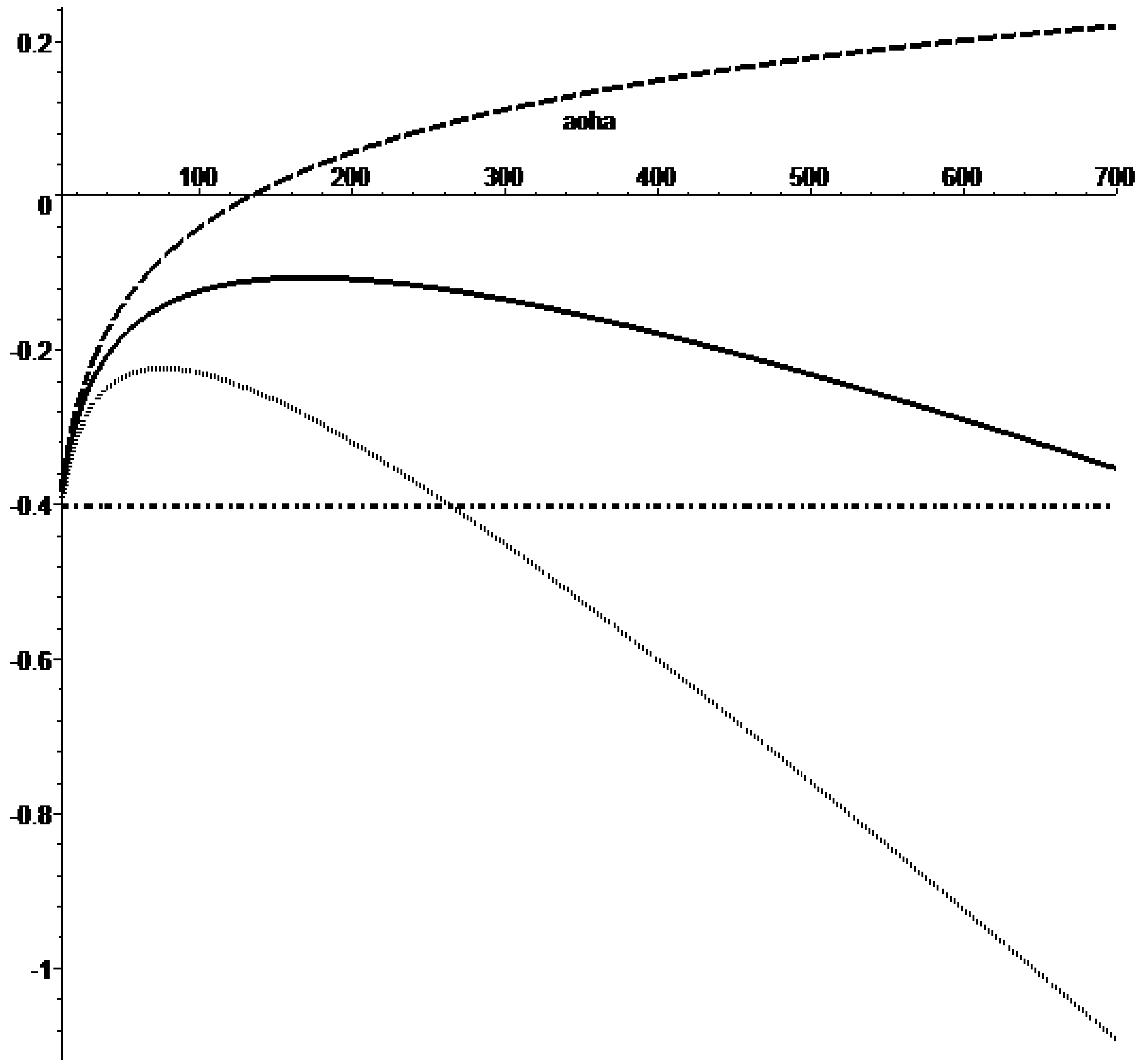


Figure 2: The graph illustrates ε_{AOHA}, the elasticity of agricultural land (AL) with respect to agricultural intensification (AOHA), plotted against AOHA. For model (1.a) (dashed-dotted line) ε_{AOHA} is constant and negative, revealing the existence of a land-sparring effect associated with intensification. For model (1.b) when corruption control (CORC) is low or moderate (i.e., CORC is evaluated at the sample minimum, dotted line, or sample mean, solid line) ε_{AOHA} is not constant but remains negative. When corruption control is high (i.e., CORC is evaluated at the sample maximum, dashed line) a Jevons paradox emerges.

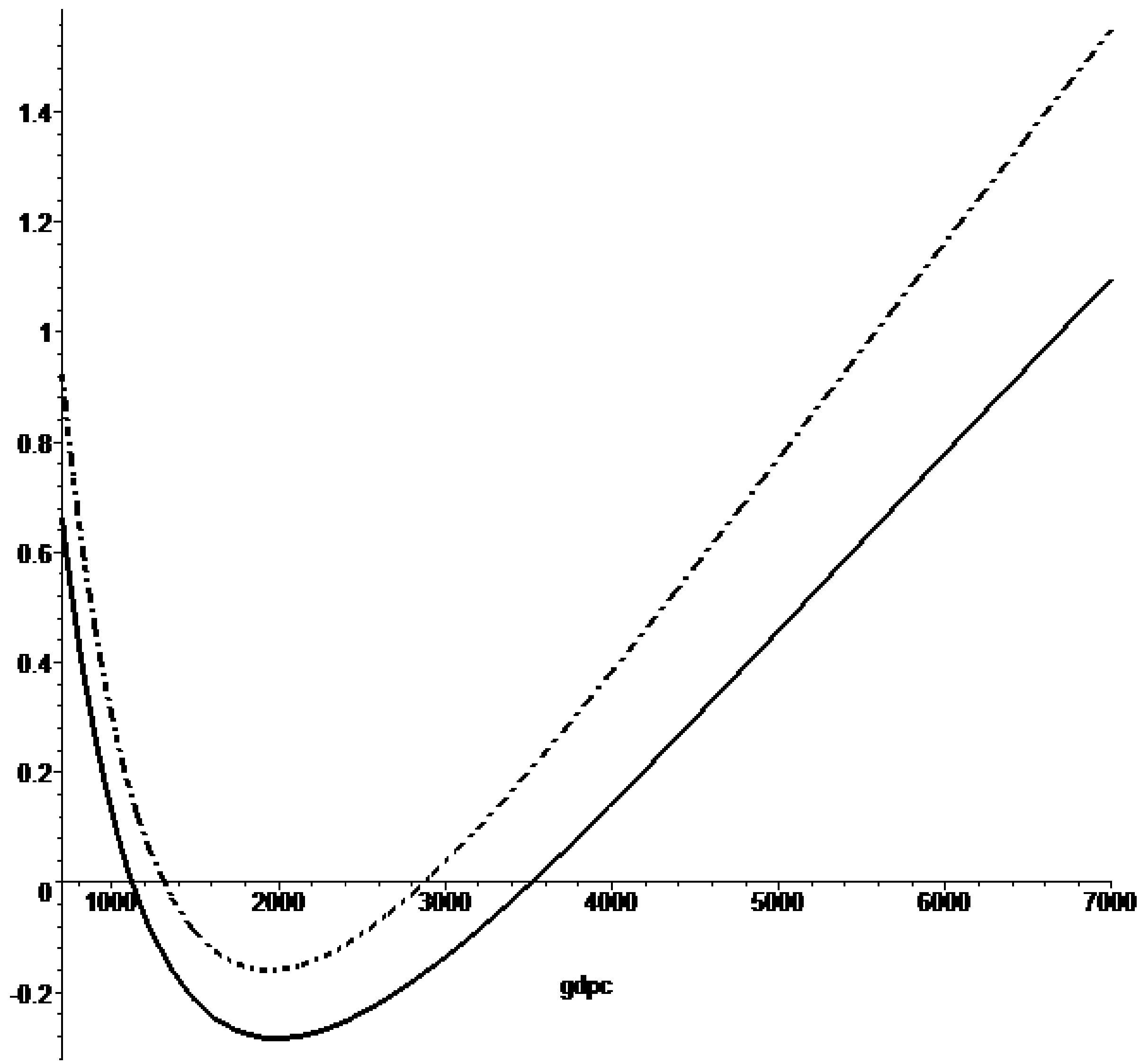


Figure 3: The graph shows the elasticity of agricultural land (AL) with respect to per-capita GDP (GDPC) for model (1.a) (dotted-dashed line) and model (1.b) (solid line).

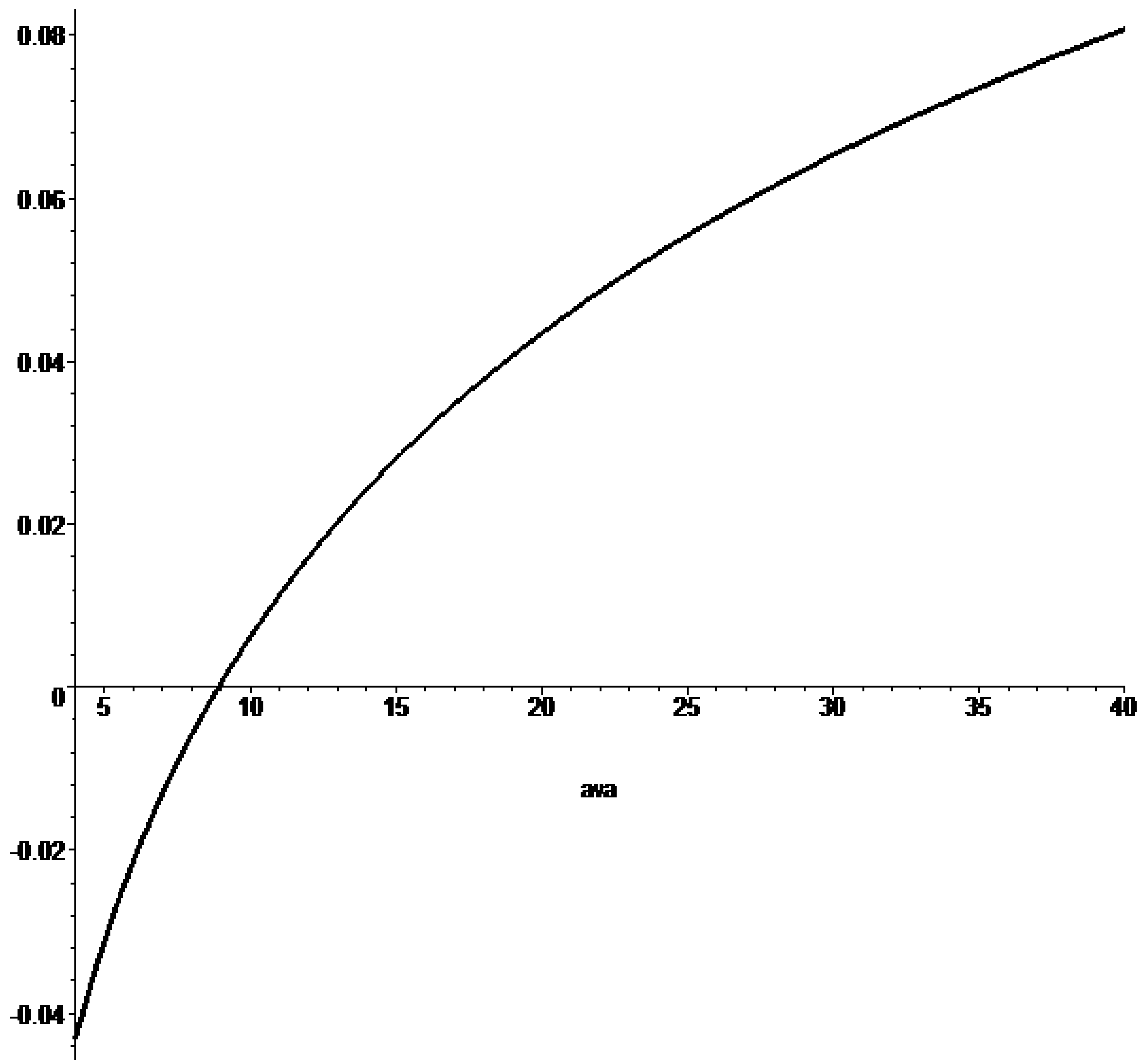


Figure 4: The graph illustrates the elasticity of agricultural land (AL) with respect to the level of service on external debt (PEDS) as a function of the share of agriculture on GDP (AVA).

Table 1: Descriptive Statistics for the sample

Variable	Description	Mean	Std	Min	Max	Cases	Source
AOHA	Value of Agricultural Output (at constant prices) per ha	156.580	108.796	16.536	641.725	217	FAO
POP	Population	37933814.9	50228349.5	2484739	187958211	217	WB
EX	Index of Agricultural Exports	63.359	46.381	4	254	217	FAO
GDPC	GDP per-capita	2525.345	1522.327	775.533	6521.484	217	WB
PEDS	Service on External Debt as % of GDP	5.314	2.513	0.923	12.111	217	WB
AVA	Agricultural Value Added as % of GDP	14.570	8.040	4.023	37.702	217	WB
CORC	Index of Corruption Control	-0.571	0.431	-1.275	-0.025	217	WB

Table 2: Regression results

Variables	Model 1.a	Model 1.b
Log(AOHA)	-0.401** (0.160)	-0.709*** (0.170)
Log^2(AOHA)	0.020 (0.015)	0.073*** (0.019)
Log(POP)	0.756§ (0.446)	1.398*** (0.452)
Log^2(POP)	0.005 (0.016)	0.005 (0.016)
Log(GDPC)	59.519*** (7.259)	49.919*** (7.292)
Log^2(GDPC)	-7.882*** (0.957)	-6.6111*** (0.962)
Log^3(GDPC)	0.347*** (0.042)	0.290*** (0.042)
Log(EX)	0.009 (0.044)	0.066 (0.044)
Log^2(EX)	0.004 (0.006)	-0.004) (0.006)
Log(PEDS)	-0.028 (0.064)	-0.118§ (0.065)
Log^2(PEDS)	0.001 (0.016)	0.003 (0.016)
Log(AVA)	-0.001 (0.050)	-0.028 (0.048)
Log(PEDS)×Log(AVA)	0.021 (0.018)	0.054** (0.019)
CORC×AOHA	-	0.001*** (0.0004)
Constant	-151.499*** (18.355)	-137.733*** (17.799)
Adj. R^2	0.997	0.997
Log-likelihood	373.046	384.199
AIC	-2.913	-3.006
Hausmann	60.75***	85.08***
LM	221.17***	262.74***
Preferred Model	2-way FE	2-way FE

Dependent variable: Log(AL); §10% s.l.; * 5% s.l.; ** 1% s.l.; *** 0.1% s.l.

Appendix – Turning points

In the main text of the article we discuss the existence of two "turning points" for the effect of per-capita GDP on agricultural land. From expressions (6.a) and (6.b) the turning points can be easily computed as

$$\tau_1 = exp\left(\frac{-2\theta_4 - \sqrt{4\theta_4^2 - 12\theta_3\theta_5}}{6\theta_5}\right) \quad \text{(A.1.a)}$$

$$\tau_2 = exp\left(\frac{-2\theta_4 + \sqrt{4\theta_4^2 - 12\theta_3\theta_5}}{6\theta_5}\right) \quad \text{(A.1.b)}$$