Agricultural Distortions, Structural Change, and Economics Growth: A Cross-Country Analysis

by

Benjamin N. Dennis
Millennium Challenge Corporation
University of the Pacific

Talan B. İşcan
Dalhousie University

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DALHOUSIE UNIVERSITY
HALIFAX, NOVA SCOTIA, CANADA
B3H 3J5

Agricultural Distortions, Structural Change, and Economic Growth: A Cross-Country Analysis

Benjamin N. Dennis[†]
Millennium Challenge Corporation,
University of the Pacific

Talan B. İşcan[‡]
Dalhousie University

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Abstract

Taxing agriculture to mobilize resources for industrialization has been a widely used development strategy. Using a novel cross-country time-series data set with direct measures of agricultural taxation, we examine how a policy bias against agriculture affects the speed of convergence in income per capita, structural change, and economic growth. We find that distortionary agricultural policies in poor economies account for the emergence of convergence clubs in our sample by significantly retarding their structural transformation and economic growth. We also identify two key channels, the subsistence consumption effect and the relative productivity effect, that account for the relatively slow structural change in poor economies with high agricultural taxes.

JEL Classification: O41, O13, O50

Keywords: taxing agriculture, structural change, convergence clubs, economic growth

[†]Economic Analysis, Department of Accountability, Millennium Challenge Corporation, 875 Fifteenth Street, NW, Washington, DC, 20005, USA. *E-mail*: DennisBN@mcc.gov. Tel.: +1.202.521.7220.

[‡]*Corresponding author*: Department of Economics, Dalhousie University, Halifax, NS, B3H 3J5, Canada. *E-mail*: tiscan@dal.ca. Tel.: +1.902.494.6994. *URL*: http://myweb.dal.ca/tiscan/.

1 Introduction

There is a wealth of evidence showing that agricultural policies differ systematically across countries: while many countries subsidize their agricultural sector, many others have consistently taxed agriculture and continue to do so using both direct and indirect methods (e.g., Anderson and Hayami, 1986). Despite the enormous significance of this issue for economic development, however, there is surprisingly little empirical analysis on the connection between such agricultural policies and economic growth performance. This paper attempts to fill this gap. Using a novel cross-country time-series data set with direct measures of agricultural taxation, we examine how a policy bias against agriculture affects the speed of convergence in income per capita, structural change, and economic growth.

In the absence of such systematic evidence, it is perhaps not surprising that there is no consensus in the literature concerning which type of agricultural policy promotes economic growth.[1] One strand in the development literature views the taxation of agriculture as a pre-requisite for mobilizing domestic savings, and regards agriculture as an abundant source of "surplus labor" that can be tapped at will to accelerate economic growth, or as a sector with limited opportunities for productivity growth.[2] According to this viewpoint, the rationale for the policy bias against agriculture in developing countries is to jump start modern economic growth.

While influential, this viewpoint has many critics who argue that, in developing countries, policies that discriminate against agriculture directly affect the incomes of the vast majority of the population, many of whom are relatively poor farmers. Since such polices push these societies closer to subsistence, they actually lead to a *reduction* in the national saving rate and, by curtailing incentives to invest, they also ultimately lead to slower productivity growth in agriculture. According to this viewpoint, the heavy taxation of agriculture in poor economies is a policy mistake that hinders economic growth.[3]

[1]Timmer (1988) and Schiff and Valdés (2002) contain excellent overviews of the relevant literature, as well as alternative viewpoints.

[2]See, for example, Lewis's (1954) classic paper and a more recent influential paper by Matsuyama (1992).

[3]See, for example, Bates (1981) and Rattsø and Torvik (2003). With reference to discriminatory agricultural policies in Sub-Saharan Africa, Mwabu and Thorbecke (2004: 32) conclude that "agriculture has continued to be the proverbial cash-cow that is unmercifully milked dry.

Given these fundamental differences, we start with a general empirical framework to examine the link between the degree of agricultural taxation and economic growth, and then extend our analysis to highlight several specific channels that explain such a link. Thus, we first ask whether there are systematic differences between discriminating and subsidizing countries in terms of their economic growth outcomes. To this end, we follow a parsimonious modeling approach in the spirit of the conditional convergence regressions found in the literature on economic growth, and we allow for clustering in income per capita driven by differences in agricultural taxes, as in the case of "convergence clubs" (Quah, 1996).

Using modern threshold regression techniques, we find that, in those countries that discriminated against their agriculture, the rate of structural change is extremely slow and the mean rate of convergence in income per capita is close to zero during our sample period. By contrast, the mean rate of convergence in the remaining group of countries in our sample is about three percent per year, accompanied by rapid structural change. In addition, we document that in countries with relatively heavy taxes on their agriculture, the total factor productivity growth rate in agriculture has also been very low. These results lead us to conclude that discriminating against agriculture has been detrimental to economic growth.

We then investigate how agricultural taxation affects economic growth in our sample using a framework that appropriately accounts for structural change and allows us to shed light on the precise transmission mechanisms at work. In this framework, agricultural taxes influence structural change primarily through two key channels, which we call the 'subsistence consumption effect' and the 'relative productivity effect.' Our econometric results show that the interaction between the subsistence nature of food consumption and agricultural taxes is a critical determinant of the speed of structural change in developing countries.

We also provide quantitative examples relevant for poor economies that illustrate the interplay across agricultural taxes, subsistence consumption, and the convergence rate of both the employment share of agriculture and income per worker to their steady states. In our calibrated economies where agricultural taxes push after-tax incomes close to subsistence, and using parameter values taken from the literature, the speed of convergence is slow (significantly less than one percent per year), which is broadly in line with our econometric estimates.

Our paper naturally overlaps with several important themes in the literature on economic development and structural change. The main organizing principle of our paper is agriculture's contribution to economic development and economic growth through structural change, and this has a long tradition in both development economics and agricultural economics; see, e.g., Tomich, Kilby, and Johnston (1995), Mundlak (2000), and Timmer (2005). We also build on literature that extensively documents the specific agricultural policies that we quantify and examine in our empirical work. Anderson and Hayami (1986) is the first systematic examination of these policies encompassing both industrial and developing countries. The volume edited by Krueger, Schiff and Valdés (1992) contains extensive documentation and a critical overview of policy biases affecting agriculture in several developing countries.[4]

Finally, our theoretical framework for structural change is closely related to the recent work by Echevarria (1997), Laitner (2000), and Kongsamut, Rebelo and Xie (2001), who examine the consequences of non-homothetic preferences for structural change, and Dennis and İşcan (2007) and Ngai and Pissarides (2007), who examine the influence of sectoral productivity growth rates on structural change. In this paper, we refer to the former as the subsistence consumption effect, and the latter as the relative productivity effect.

The rest of our paper is organized as follows. Section 2 sets the stage by demonstrating the empirical regularities we find in our data set using bivariate correlations, and then presents the estimates of multivariate convergence regressions that link agricultural taxes to economic growth and structural change. Section 3 presents a conceptual framework that demonstrates the impact of agricultural taxes on structural change. Section 4 presents the estimates of regressions that link structural change to its key drivers, including agricultural taxation. Section 5 concludes.

[4]For recent discussions of discrimination against agriculture, see Purcell and Gulati (1993) for India, Takeuchi and Hagino (1998) for the Philippines, and Akiyama et al. (2003) for Africa, and for a historical perspective, see Lindert (1991). Anderson (2006) provides a recent appraisal of this literature and discusses the welfare consequences of both agricultural taxes and subsidies."

2 Agricultural taxes and economic growth

The key to our analysis and the novel variable of our cross-country time-series data set is a measure of the taxation (subsidy) policies that are biased against (in favor of) agricultural producers. We follow the lead of Anderson and Hayami (1986) and others, and use the agricultural *nominal rate of protection* (henceforth, NRP) as a measure of agricultural distortions (taxes and subsidies). In the case of agricultural taxes, the NRP is negative and domestic farmers receive less than the international price per unit of output they produce. In the case of subsidies, the NRP is positive and domestic producers receive more than the international price per unit of output. We present a detailed discussion of this and other variables we use in this study in Appendix A. The data set covers 47 countries (listed in the data appendix) and two periods from 1960 to 1972, and from 1976 to 1984.

2.1 Empirical regularities

Our analysis synthesizes five empirical regularities we observe in our cross-country data: (i) the prevalence and persistence of the NRP, (ii) the positive association between the NRP and GDP per capita, (iii) the positive association between the NRP and economic growth, (iv) the positive association between the NRP and agricultural total factor productivity (TFP) growth, and (v) the positive association between the rate of reallocation of labor out of agriculture (i.e., structural change) and economic growth. We discuss each of these in turn, and then present a formal multivariate analysis linking the NRP to economic growth.

NRP.—It is well-known that agriculture has been a heavily distorted sector in both rich and poor countries. Figure 1 shows that the majority of the countries in our sample use agricultural policies that lead to either agricultural taxes or subsidies, and that there is a wide cross-national variation in the NRP. The NRP is also highly persistent in each of these countries. For example, very few discriminating countries during the period 1960–1972 turned into agricultural protectionists during the period 1976–1984, and only one protectionist country from 1960 to 1972 began discriminating against agriculture in the latter period.

GDP per capita.—At least since the seminal work of Anderson and Hayami (1986), it is also well-known that policies tend to discriminate against agriculture in developing countries and tend to protect agriculture in industrial countries. This relationship holds remarkably well in our larger sample of countries as well. Figure 2, panel (a), shows the positive relationship between real GDP per capita and the NRP.[5]

Economic growth.—Here, we document a less well-known fact about agricultural distortions: the strong negative relationship between agricultural taxation and economic growth. Figure 2, panel (b), relates the NRP to the growth rate of annualized real GDP per capita, and shows that the bivariate relation between the NRP and economic growth is negative. This is the key empirical regularity we explore in detail below.

TFP.—The fourth empirical regularity we document in figure 3 is also less well-known, but not so surprising in light of what we have documented so far: there exists a strong negative correlation between agricultural taxation and agricultural TFP growth.[6] These striking patterns are suggestive of a causal relation whereby the relatively poor aggregate growth performance of some developing countries is in part due to those economic policies that were heavily biased against their large agricultural sectors.

Structural change.—The fifth empirical regularity we capture in our analysis concerns structural change. At least since Kuznets (1966), economists have extensively documented the strong negative association between real income per capita and the share of labor in agriculture (see also Chenery and Syrquin, 1975). Figure 4 shows that these patterns also hold in our data set. The common interpretation of these patterns is that countries that have become rich did so by reallocating labor out of agriculture. However, there is much less systematic evidence about how such a reallocation of labor has been possible for some countries, but not for others. As we show in section 4, differences across countries in terms of the agricultural policies identify go a long way in explaining such differences.

[5]This relationship is clearly not mediated through openness to international trade. Figure 1 shows the relation between the NRP and the (subjective) classification of "outward orientation" published by the World Bank (1987). Evidently, both outward- and inward-oriented economies distort domestic agricultural prices, and economic policies that result in agricultural distortions are not highly correlated with broader trade policy choices: see also section 2.4.

[6]See Fulginiti and Perrin (1993, 1999) for further discussion of this relationship, but for a smaller sample of countries. Restuccia et al. (2006) examine the quantitative importance of barriers to technology adoption in agriculture in accounting for cross-country variation in income levels.

2.2 Convergence regressions

We now turn to a more systematic analysis of our data and investigate the relation between the NRP and economic growth within a multivariate framework. Our framework at this stage does not take a stance on a specific link between agricultural taxation and economic growth. Rather it exploits a direct implication of all the strands in the theoretical literature on the nexus of agriculture and economic growth, that agricultural policy is a key determining factor of cross-national growth behavior, both in terms of income levels and convergence rates.

The conventional approach to convergence (e.g., Barro and Sala-i Martin, 1995, chapter 11) typically estimates the average convergence rate in a cross-section of countries, and thus estimates the convergence rate for a representative country, after accounting for steady-state income level differences. By contrast, the distinct empirical methodologies proposed by Easterly (1994), Durlauf and Johnson (1995), Quah (1996), and Canova (2004) start from the premise that national per capita income levels may display "convergence clubs," and therefore allow for variations in observed growth behavior across these groups of countries. In our convergence analysis, we combine these two complementary strands in the literature and examine how cross-national differences in agricultural taxation affect rates of convergence after accounting for those variables that lead to differences in steady-state income levels.[7]

Testing for differences in cross-national convergence rates based on agricultural taxation amounts to testing for different poles of attraction in our data, and for this purpose we use the threshold estimation and inference methods developed by Hansen (2000). This method involves assessing whether convergence rates across countries differ based on an observable variable, in our case, the nominal rate of protection, and as such it nests the standard convergence regression as a special case.

In particular, consider the following version of the conditional convergence regression

[7]Given that we have less than thirty years of observations in our cross-section of countries, we cannot statistically distinguish between long-run growth and level effects due to agricultural taxation. However, since many of the countries in our sample have not had completed their structural transformation, we design our empirical analysis to appropriately account for growth effects, even if these are transitory.

model:

$$T^{-1}(\ln Y_{Tj} - \ln Y_{0j}) = a + a\mathbf{1}(\tau) - b \ln Y_{0j} - b \ln Y_{0j}(\tau) + A\mathbf{X}_j + A\mathbf{X}_j(\tau) + u_{Tj}. \quad (1)$$

where $j = 1, \ldots, J$ is a country index, $\mathbf{1}$ is a column vector of ones, $b = \frac{(1-e^{-\beta T})}{T}$, where $\beta > 0$ is the rate of convergence, and $Y_j(\tau) = Y_j d_j(\tau)$ with the dummy variable $d_j(\tau) = \{\tau_{Aj} < \tau\}$, where $\{\cdot\}$ is the indicator function (similarly for $\mathbf{1}(\tau)$ and $\mathbf{X}_j(\tau)$). A is a vector of parameters, and $\mathbf{X}_j$ is a matrix of conditioning variables (see below). To accommodate the possibility of convergence clubs, the model allows the regression parameters to differ depending on the magnitude of the threshold variable τ_{Aj} relative to the threshold parameter τ. The objective is to estimate the threshold parameter together with the remaining parameters of the model. If the threshold parameter is not statistically significant, the model collapses into the familiar conditional convergence regression whereby only initial income and a set of conditioning variables matter for observed growth.

Because there is a close association in our data between income levels and structural change, we simultaneously test for both convergence in income per worker and the non-agricultural sector's share of employment. Consequently, we estimate the regression model (1) using the growth rates of both real GDP per capita in PPP prices (rgdp) and the non-agricultural employment share (L_M) as the dependent variables. We interpret the rate of change in L_M as the speed of structural change—as commonly done in this literature.

Our conditioning variables in the convergence regression (1) are taken directly from the conditional convergence literature (see, e.g., Durlauf and Johnson, 1995) and include the natural logarithms of the investment to GDP ratio, the gross enrollment ratio for primary education, and the growth rate of population (plus 0.05 to account for the depreciation rate and the growth rate of labor augmenting technology). The employment share convergence regressions include initial income per capita as well.

We pool the data and use the average growth rates from 1960 to 1970 and from 1970 to 1980 for our sample. For all conditioning variables, we use values corresponding to the initial year (i.e., 1960 or 1970) except the growth rate of the population, which is computed as the average growth rate over the corresponding period. We do this to ensure that our conditioning variables are predetermined (relative to the agricultural taxation), and therefore are not jointly determined by the NRP. This issue is particularly relevant for the rate of investment, which would be highly influenced by contemporaneous agricultural taxation through its influence on domestic savings.

Table 1 presents descriptive statistics for some of the variables used in the empirical analysis. We group the variables under three broad headings: nominal rate of protection, income and economic structure, and other measures of distortions (discussed below). Under "income and economic structure", we include real GDP per capita, and those variables that are related to structural change (such as the sectoral distribution of income and employment). The mean NRP is negative in our sample but close to zero in both periods, and it varies considerably across countries. Countries that tax and those that subsidize their agriculture are both well represented in the sample.[8]

2.3 Convergence results

As mentioned above, the first step in our empirical analysis is to establish whether there is indeed evidence for threshold effects driven by agricultural taxation. Hansen (1996) proposes a heteroskedasticity-consistent Lagrange multiplier test for a threshold.[9] The test results suggest that the NRP is a strong determinant of threshold effects in our data set. We find that for per capita income convergence, the null of "no threshold effect" has a p-value of 0.066, and for employment share convergence the p-value is 0.001 leading us in both cases to strongly reject the null of no threshold effect in growth rates due to the NRP.

Given that there is evidence of agricultural-tax-driven threshold effects, the next step is to estimate the value of the NRP that corresponds to the threshold parameter. We estimate this threshold τ using the procedure discussed in Hansen (2000). Intuitively, τ is the value that minimizes a normalized likelihood ratio function. Figure 5, panel (a)

[8]We do not claim that our sample is representative of all countries in the world. For example, our analysis is not applicable to land-scarce city-states, unless these are modeled as part of a broader geographical entity with an agricultural hinterland.

[9]As discussed by Hansen (1996), the threshold τ is not identified under the null of no threshold effect. Hansen proposes a bootstrap method that delivers asymptotically correct p-values. The p-values we report are based on this procedure with 1,000 bootstrap replications.

displays the likelihood ratio sequence as a function of the threshold in the NRP, $LR_n(\tau)$, when the dependent variable is per capita income, and panel (b) displays the $LR_n(\tau)$ when the dependent variable is L_M. For the convergence in real GDP per capita, the point estimate of the threshold is NRP $= -4.60$, and the asymptotic 95% confidence interval is $[-8.00, 34.33]$, suggesting that there is substantial uncertainty about the precise value of the threshold when rgdp is used as the convergence criterion. For the convergence in L_M, the point estimate of the threshold is NRP $= -7.00$, and the asymptotic 95% confidence interval is $[-9.00, 0.00]$, suggesting that, in this case, the threshold is precisely estimated. (Complete estimation results are available from the authors upon request.)

Using these threshold parameter estimates, we also compute the convergence rates corresponding to each of these two convergence clubs and find the differences in implied convergence rates in per capita income economically significant. For the group of countries with NRP< -4.6, we cannot reject that the implied convergence rate equals zero (i.e, $\hat{b} = -0.0021$ with standard error 0.0038). For the group of countries with NRP> -4.6, however, the convergence rate is slightly under 3 percent per year and the coefficient on initial income per capita is precisely estimated (i.e, $\hat{b} = -0.025$ with standard error 0.0048). Similarly, we find that the implied convergence rates in employment shares are also significantly different across the two convergence clubs we identify. For the group of countries with NRP< -7.0, the coefficient on the initial employment share is substantially smaller in absolute value compared to the coefficient for the group of countries with NRP> -7.0.[10]

Ideally, we would like to test whether there are multiple threshold variables that lead to differences in convergence rates across countries. Unfortunately, the distribution theory for inference has only been developed for a single threshold variable thus far, and we cannot run a test across potentially competing threshold variables or jointly estimate these threshold parameters.[11] The threshold variable that has attracted the most attention in

[10]In the employment share convergence regressions, the estimated coefficients on the initial employment share and their standard errors (in parentheses) are as follows. For the group of countries with NRP< -7.0 $\hat{b} = 0.286$ (0.5294), and for NRP> -7.0 $\hat{b} = -0.723$ (0.367), corresponding to striking differences in the rate of structural change.

[11]There are two other related issues that we cannot satisfactorily address given data limitations. First, there may be multiple thresholds and thus more than two convergence clubs. Second, there may be (additional) clusters of income per capita within each convergence club. Both of these are elegantly addressed

previous research is initial per capita income (e.g., Durlauf and Johnson, 1995), which we also find to be a valid threshold variable for both per capita income and employment share convergence. (These results are available from the authors upon request.) In our context, this finding has a natural interpretation: low per capita income is associated with a high expenditure share of subsistence food consumption, and subsistence consumption amplifies the influence of agricultural taxes on economic growth (see, e.g., Steger 2000).

To further pursue the possibility that combined influences of initial income and taxes would be stronger than that of taxes alone, we use a term that interacts log real GDP per capita and the NRP (where the interaction term is labeled, "GDPNRP") as a threshold variable. Figure 5 presents the likelihood ratio sequence as a function of the threshold in GDPNRP, $LR_n(\tau)$, when the dependent variable is the average growth rate of per capita income (panel (c)), and when the dependent variable is the average growth rate of L_M (panel (d)). In both cases, there is statistically significant evidence for threshold effects based on GDPNPR (p-values for the null of no threshold effect are 0.050 and 0.001, respectively). Also, in both cases, consistent with the idea that agricultural taxation pushes poor economies closer to subsistence and amplifies differences in convergence rates, the thresholds correspond to a combination of a relatively low income per capita and high taxation of agriculture.

2.4 Sensitivity analysis

It is natural to ask whether, and, if so, how the NRP relates to other measures of distortions used in the literature, including the parallel market premium, and taxation and other forms of discrimination against investment goods, particularly imported ones. Appendix A discusses these other measures of distortions in further detail and shows that the agricultural taxes we use are not highly correlated with other measures of price distortions.[12] Of these measures, the NRP is unique in its focus on sectoral distortions that

by Canova (2004), who extends Hansen's (2000) approach to multiple thresholds (but maintains the unique threshold variable assumption). Unfortunately, Canova's Bayesian techniques are data intensive, and are not feasible with our data set. We do, however, provide an informal test of the second possibility in section 2.4.

[12] We also considered the bivariate correlations between the NRP and two measures of international trade connectedness: exports plus imports divided by GDP based on current prices, all based on national

directly affect domestically produced goods, and in that it treats taxes and subsidies symmetrically.

In any case, we find that the NRP continues to have both strong level and convergence club effects when we use other distortion measures as control variables in equation (1). These results are presented in table 2 for convergence in GDP per capita and in table 3 for convergence in the employment share of non-agriculture. Specifically, we consider four regression specifications (see table 2). Regression 1 is the baseline model in equation (1). Regression 2 includes the NRP separately as an independent variable. Regressions 3 and 4 further introduce, respectively, the parallel market premium and the relative price of investment into regression 2 as additional controls. Table 2, panel a, shows estimation results when no threshold effects are imposed. In all these regressions, when we include the NRP in the regression model separately as an independent variable, its coefficient is statistically different from zero.[13] In general, we also find strong evidence against the null hypothesis of no threshold effects while controlling for these other variables. [14]

Furthermore, to check for the existence of clusters of income driven by NRP within each convergence club, we split the data into two groups using the NRP as the threshold variable. We find that the NRP (and other distortions) has no significant effect on the level of steady-state income *within* each group—although the coefficient on initial income is always significant. Using the same methodology, our findings are even stronger for the impact of the NRP on the growth rate of non-agriculture's employment share; see table 3).[15] In fact, even when we (incorrectly) ignore the threshold effects, the coefficient

income accounts, and exports plus imports, both in current U.S. dollars (and both from the balance of payments accounts) divided by GDP in current international prices. These correlations were also significantly low, and are not reported to conserve space.

[13]The relative price of investment appears to independently influence growth and serves as a complementary measure of sectoral distortions. By contrast, the coefficient on the parallel market premium is not statistically significant in our sample, although we do not rule out the possible influence of this variable on growth in a broader sample of countries.

[14]When we allow for threshold effects due to differences across countries in the NRP, the p-values for the null of no threshold effects are very low, with the possible exception of the regression model which simultaneously controls for the NRP and the relative price of investment in the growth regression. Even in that case, however, the p-value is about 40 percent, and the estimated NRP threshold value is about 17 percent.

[15]The reported parameter estimates in table 3, panel a, are from those regressions, which impose no threshold effects. In all cases, there is overwhelming evidence against the null of no threshold effects driven by the NRP, as the p-values for the null hypothesis are always less than 0.006. And, in all these

on the NRP is still significant, whereas none of the two other measures of distortions we consider are statistically different from zero. We thus conclude that the NRP contains information about convergence rates that is distinct from that contained in previous studies.[16]

3 Inspecting the channels

In the context of structural change, agricultural taxation affects the speed of convergence through several channels. To identify and shed light on the empirical relevance of the precise channels at work, we consider a two-sector model with agriculture and non-agriculture. The two prominent channels that can account for structural change are non-homothetic preferences (as in Kongsamut et al. 2001), and differential productivity growth across sectors (as in Dennis and İşcan 2007, and Ngai and Pissarides 2007). The model encompasses both of these channels, illustrates how agricultural taxation influences the sectoral allocation of labor, and has testable empirical implications.[17]

3.1 A basic framework

There are two sectors: agriculture, A, and non-agriculture ("manufacturing"), M. We adopt a representative agent setup.

Preferences.—Consumption is the only determinant of instantaneous utility. At time t, the composite consumption good C is given by an CES-aggregator function

$$C_t = \left[\eta^{1/\nu} C_{Mt}^{(\nu-1)/\nu} + (1-\eta)^{1/\nu} \left(C_{At} - \gamma_A\right)^{(\nu-1)/\nu}\right]^{\nu/(\nu-1)}, \tag{2}$$

cases, the threshold value for the NRP is consistently estimated to be between −9 and −6 percent.

[16] While we do not find a close association between heavy agricultural taxation and trade policies that distort other prices, a related and equally important issue is whether taxing countries also had poor "institutions" as distinct from those that set their agricultural policies. We are unable to judge with any confidence whether those countries that heavily tax their agriculture also have higher levels of corruption, inadequate rule of law, and poor corporate and political governance proxies because most of these proxies start in 1984 and do not overlap significantly with our sample.

[17] In the model, non-homothetic preferences arise because of subsistence food consumption, which is especially relevant for poor developing countries. See Clark and Haswell (1964) on the interactions between food requirements and organization of economic activity, and Fogel (2004) on the interplay between nutrition, health and economic development and growth.

where C_M is the consumption of the non-agricultural good, C_A is the consumption of the agricultural good, $\gamma_A \geq 0$ is the subsistence level of food consumption with $C_A > \gamma_A$, η is a parameter between 0 and 1, and $\nu > 0$ is the elasticity of substitution between food consumption (net of subsistence) and non-food consumption. Empirical estimates of ν are significantly less than one, implying that agricultural and non-agricultural goods are gross complements. When $\gamma_A > 0$, preferences are non-homothetic.

Production.—Output, Y, in each sector is given by

$$Y_{At} = B_A K_{At}^{\alpha} (Z_{At} L_{At})^{1-\alpha}, \tag{3}$$

$$Y_{Mt} = B_M K_{Mt}^{\alpha} (Z_{Mt} L_{Mt})^{1-\alpha}, \tag{4}$$

where B_i $(i = A, M)$ is an efficiency parameter, K_i is the sectoral capital stock, $0 < \alpha < 1$ is the income share of capital, Z_i is sectoral labor augmenting technology, and L_i is sectoral labor services.

Agricultural taxes.—We consider the NRP as a flat tax rate τ_A on agricultural production.[18] This formulation is consistent with the literature that examines the influence of economywide distortions on economic growth (e.g., Easterly 1994).

Aside from their direct impact on income, as suggested by figure 3, agricultural taxes may also influence the adoption of more productive—but possibly more risky—technologies and slow down the rate of agricultural total factor productivity growth.[19] We allow for this possibility, and let agricultural labor augmenting technology depend on τ_A. Thus, we write $z_t(\tau_A)$ where $z_t = Z_{Mt}/Z_{At}$.

Resource constraints.—Factors of production are constrained at the aggregate level by

$$K_{At} + K_{Mt} \leq K_t \quad \text{and} \quad L_{At} + L_{Mt} \leq 1. \tag{5}$$

Total employment is normalized to one, and all aggregate variables can be interpreted in per worker terms.

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[18] Of course, when $\tau_A > 0$, it must satisfy a "no starvation" constraint, $(1 - \tau_A) B_A K_t^{\alpha} Z_{At}^{1-\alpha} > \gamma_A$.

[19] Under uncertainty, the non-homothetic preferences in the model imply a coefficient of relative risk aversion that declines with income. So, it is plausible to think that agricultural taxes would bias the farmers' choices towards traditional technologies with both lower risk and productivity. Also, agricultural marketing boards that control agricultural prices tend to have other functions as providers of seed and fertilizers, and as such may be barriers to technology adoption.

The agricultural good can only be consumed, C_A, but the manufactured good can either be consumed, C_M, or invested in physical capital:

$$C_{At} = (1-\tau_A)B_A K_{At}^{\alpha}(Z_{At}L_{At})^{1-\alpha}, \tag{6}$$

$$I \equiv \frac{dK_t}{dt} + \delta K_t = B_M K_{Mt}^{\alpha}(Z_{Mt}L_{Mt})^{1-\alpha} - C_{Mt}, \tag{7}$$

where $0 \leq \delta < 1$ is the depreciation rate.

Production efficiency.—There is perfect factor mobility across sectors. This implies the equality of marginal rates of transformation across sectors, or

$$\frac{K_{At}}{Z_{At}L_{At}} = \frac{K_{Mt}}{Z_{Mt}L_{Mt}}. \tag{8}$$

Intratemporal consumption allocations.—The following first-order condition characterizes the optimal allocations of A- and M-goods:

$$\left(\frac{1-\eta}{\eta}\right)\left(\frac{C_{Mt}}{C_{At}-\gamma_A}\right) = P_{At}^{\nu}. \tag{9}$$

Prices.—Normalize the price of the M-good to 1, and define P_A as the (relative) price of the A-good. Product markets are competitive, so we have

$$P_{At} = \frac{bz_t(\tau_A)}{1-\tau_A}, \tag{10}$$

where $b = B_M/B_A$.

3.2 Sectoral allocation of labor

With an eye toward the empirical implementation of the model, we use equations (3)–(10), and express the share of labor in the M-sector as

$$L_{Mt} = [1+p(z_t)]^{-1} \times [p(z_t)s_{Mt} + (1-s_{At})], \tag{11}$$

where the *subsistence consumption* effect, which captures the influence of non-homothetic preferences, is

$$s_{At} = \frac{\gamma_A}{(1-\tau_A)Y_{At}/L_{At}}, \tag{12}$$

the *relative productivity* effect on the employment share of non-agriculture is

$$p(z_t) = \left(\frac{1-\eta}{\eta}\right)\left(\frac{b}{1-\tau_A}\right)^{1-\nu} z_t(\tau_A)^{1-\nu}, \tag{13}$$

and the *capital accumulation* effect, or the influence of the rate of (gross) investment on the employment share of non-agriculture, is

$$s_{Mt} = \frac{I_t}{Y_{Mt}/L_{Mt}}. \tag{14}$$

In the above expressions, the subsistence consumption effect captures the fact that income elasticity of agricultural goods is less than one, and as agricultural productivity increases, agriculture's share of employment declines. The relative productivity effect captures the empirically relevant gross complementarity $\nu < 1$ between agricultural and non-agricultural goods. When the elasticity of substitution between agricultural and non-agricultural goods is less than one (complementarity), higher productivity growth in agriculture relative to that of non-agriculture leads to a decrease in agriculture's share of employment, and vice versa. The capital accumulation effect captures the fact that only non-agriculture produces investment goods, and capital deepening leads to an increase in non-agriculture's share of employment.

Finally, interacting with these three channels is the effect of agricultural taxation on structural change. Agricultural taxation influences all three channels and slows down structural transformation, especially in poor countries. First, agricultural taxes burden farmers close to subsistence food production, both directly as claims on farmers' current incomes and indirectly through their effect on the choice of technology. Both factors tend to lead to a reduced supply of food at a higher unit cost. To compensate for this tax-induced short-fall in food production, agriculture ends up accounting for a relatively higher share of total employment (the subsistence consumption effect). Second, agricultural taxes influence the domestic terms of trade, again both directly since a higher tax is similar to a reduction in agricultural productivity, and indirectly through the choice of technology. Given gross complementarity between agricultural and non-agricultural goods, both lead to a slower reallocation of labor out of agriculture (the relative productivity effect). Third, agricultural taxes reduce disposable income and thereby savings, again especially in poor countries where subsistence food consumption accounts for a large share of disposable income. Consequently, lower savings reduce the demand for investment goods, reducing in turn the demand for labor in non-agriculture (the capital accumulation effect).

It is easy to see that for poor countries the influence of agricultural taxes on structural change is primarily mediated through the subsistence consumption effect. Intuitively, given that agriculture produces a hard-to-substitute subsistence good, poor economies find it extremely difficult to reallocate resources out of agriculture, especially when taxes simultaneously retard agricultural TFP growth. Consequently, the combination of low productivity and a high share of food in total consumption expenditures renders poor countries particularly susceptible to the adverse effects of agricultural taxes.[20]

In the next section, we empirically examine the impact of agricultural taxation on these principle drivers of structural change.

4 Agricultural taxes and structural change

Our empirical analysis relates sectoral reallocation of labor out of agriculture $\hat{L}_{Mt}$ to the three principle drivers of structural change identified in the previous section: the capital accumulation effect (as captured by the investment rate), the subsistence consumption effect (as captured by the ratio of agricultural to non-agricultural employment at the beginning of the sample period, $(1 - L_{M0})/L_{M0}$), and the relative productivity effect (appendix B presents the derivation of the regression model starting from equation (11)):[21]

$$\hat{L}_{Mt} = b_0 + b_1 \text{ Investment rate}(\tau_A)_t + b_2 \left(\frac{1 - L_{M0}}{L_{M0}}\right) + b_3 \text{ Relative productivity}(\tau_A)_0 + \varepsilon_t. \tag{15}$$

The regression model assigns specific weights to each of these effects, and in our empirical analysis we investigate whether the estimated coefficients are consistent with the framework developed in the previous section. More importantly for our purposes, in the regression model, investment rate and relative productivity effects are directly influenced

[20]Our basic multi-sector model can be easily extended to examine its intertemporal implications. In the next section, we illustrate the speed of convergence in the neighborhood of steady-state income per capita for such a model.

[21]By considering the changes in the sectoral allocation of non-agricultural labor, rather than its level, we account for the potential influence of country-specific fixed effects, such as geography and natural endowments, on structural change.

by agricultural taxation. Therefore, to assess the significance of agricultural taxation on structural change, we first estimate equation (15) using investment rates and relative productivity levels but without making any adjustments for agricultural taxation. This, under our maintained hypothesis, yields a misspecified model. We also estimate the model after constructing our variables by properly adjusting them for the presence of agricultural taxation (see appendix B for details). We then examine whether the model with agricultural taxation performs better than the model without agricultural taxes using several statistical and economic criteria.

4.1 Regression results

Table 4, panel (a), reports for regression equation (15) the parameter estimates for pooled data, and for the models with and without an adjustment for agricultural taxes. Although qualitatively both sets of estimates are similar, controlling for agricultural taxation leads to important differences that are relevant for our analysis. First, accounting for agricultural taxes leads to a considerable improvement in the in-sample goodness-of-fit as shown by an increase in the adjusted R^2. Second, with agricultural taxes the coefficient estimates on employment shares increase significantly. According to our theoretical framework, the coefficient on employment shares in part captures the influence of subsistence consumption on structural change (see appendix B). Consequently, such an increase imparts a larger role to subsistence consumption effect. This result is especially relevant within our context, because, as we will discuss in detail below, the interaction between the subsistence consumption and agricultural taxation is critical for both slow structural change and slow convergence in income per capita in poor economies.

In addition, the model makes the prediction that the coefficient on the investment rate is one. The point estimates are below one for the model without distortions, and above one for the model with distortions—although in both cases these coefficient estimates are not statistically different from zero at conventional levels.

We also calculated Vuong's (1989) statistic to perform a likelihood ratio test for model selection and non-nested hypotheses. Note that models with and without agricultural taxes are (partially) non-nested. Under the null hypothesis that the two models are equivalent, Vuong's statistic has a standard normal distribution. When we test the baseline model in column 1 against the model in column 3 (table 4, panel a), Vuong's statistic gives -0.333 with a p-value $= 0.63$. Although the results are more favorable for the model that accounts for agricultural taxes, statistically we cannot discriminate between the two models.

Because the agricultural taxes influence $\hat{L}_M$ entirely through the relative productivity, capital accumulation and subsistence consumption effects, the regression model does not call for an independent influence of the NRP on structural change. Indeed, the results in table 4 show that including the NRP as an independent variable does not improve the statistical fit of the model, and the coefficient estimate on the NRP is economically small and not statistically different from zero. The results are the same when we test for nonlinear effects by including the squared NRP instead of simply the level of the NRP. We thus conclude that our regression equations appropriately capture the links between agricultural taxes and structural change.

There is also an indirect but interesting way to assess the plausibility of our estimates. Relative productivity growth in favor of the agricultural sector manifests itself through lower agricultural prices, and ultimately leads to the reallocation of labor across sectors; see equations (13). In our regression model, the coefficient on the relative productivity effect, b_3, informs us about this channel because it reflects the joint influences of gross complementarity between agricultural and non-agricultural goods and relative agricultural productivity growth. Specifically, given that the elasticity of substitution between agricultural and nonagricultural goods is less than one ($\nu < 1$), a positive coefficient on relative productivity $\hat{b}_3 > 0$ would suggest that, in our sample, agricultural productivity growth has exceeded that of the non-agricultural sector. Our estimate of this coefficient is positive and it is precisely estimated. Consequently, our estimates imply that, on average, the contribution of relative productivity growth to structural change has been positive.

Two pieces of independent evidence are consistent with this finding. First, the relative price of food items declined from the 1950s through the late 1980s in world markets (Anderson and Hayami, 1986) and individual countries (Mundlak, 2000), suggesting that relative productivity growth has been in favor of agriculture worldwide. Second, over the same period, agricultural labor and total factor productivity growth rates have consistently exceeded their non-agricultural counterparts in most countries (Mundlak, 2000; Martin and Mitra, 2001). This superior performance in agriculture in conjunction with gross complementarity between agricultural and non-agricultural goods has therefore on average been an important contributing factor to the reallocation of labor out of agriculture worldwide—although there has clearly been massive cross-country variation in agricultural TFP growth rates (figure 3).

4.2 Sensitivity analysis

To check the sensitivity of our results to pooling the data, we also estimated equation (15) for the decades 1970 and 1980 separately. Table 4, panel (b), shows the estimation results. We display only those results that properly adjust for agricultural taxation because, as in the case of pooled estimates, specification tests are more favorable to the model with taxes. The estimates suggest that the empirical model explains the data reasonably well for the 1970s (with an adjusted R^2 of 71 percent), but its performance declines for the 1980s.

We performed additional sensitivity analysis and found that our baseline regression results are remarkably robust.[22] First, we allowed the coefficient on the employment shares, b_2, which captures the subsistence consumption affect to vary across decades (while keeping other parameters constant across periods), but this did not affect our results. The coefficient estimate on the employment share is remarkably stable across our sample period, and we cannot reject the equality restriction. This suggests that the influence of subsistence consumption effect on structural change was stable over the 1970s and 1980s. Since poor countries tend have a larger share of income devoted to subsistence consumption, and since during these decades economic growth in poor countries was particularly disappointing, we also find the stability of this parameter estimate economically plausible.

Second, the theoretical model permits an interaction term that involves the rate of investment and the relative productivity growth rate. We allowed for this possibility by multiplying the share of investment in nonagricultural output by the NRP (divided by 100), and incorporated this interaction term into the baseline specification with agricul-

[22]The detailed results are available from the authors upon request.

tural taxes. For both the pooled and cross-section regression models with the interaction term, the estimation results were very similar to those from the baseline model. However, Vuong's statistic decidedly favored the baseline model over these alternatives.[23]

4.3 An illustrative example

The econometric results above suggest that agricultural taxes had a significant influence on structural change primarily through two channels; subsistence consumption and relative productivity.[24] We now ask whether these channels can also account for the convergence rates we found in section 2. In particular, the empirical results show that for the group of countries that tax their agriculture, the convergence rate of income per capita is significantly lower than the rest of the countries in our sample, and we ask whether a plausibly calibrated version of the theoretical model of structural change in section 3 can generate extremely slow convergence to steady-state income per capita.

To this end, we calibrate the model parameters and set the productivity parameters B_A, and B_M to unity (without loss of generality), the weight of non-agricultural goods to 0.85 (to match an equivalent steady-state share of non-agriculture), and the elasticity of substitution between agricultural and non-agricultural goods to 0.1. The data in figure 3 suggest practically zero agricultural TFP growth in countries with heavy agricultural taxes. It also is well-known that throughout the 1970s and 1980s the TFP growth rate in the stagnant economies has been very low or even negative (e.g., Ndulu and O'Connell 1999, table 1). These suggest a TFP growth rate of about zero, with the implication that, given initial conditions, all structural change is driven by subsistence consumption and capital accumulation effects, as income per capita increases to its steady state due to capital accumulation.[25]

[23]We have also checked the sensitivity of our results to the inclusion of the enrollment rate in primary school (proxy for human capital) in our structural change regressions. The results (available upon request) show that, once we control for agricultural taxation and the three principle channels we identified above, this control contains no additional information about structural change.

[24]Although the regression estimates are unable to identify the precise influence of agricultural taxation on the capital accumulation effect, this effect may still contribute to structural change.

[25]We embedded the intratemporal aspects of the model of section 3 into an otherwise standard Ramsey growth model, and log-linearized the dynamic system around its steady state (e.g., Barro and Sala-i-Martin 1995, pp. 87–88), and calibrated the benchmark model using, for the remaining parameters, standard parameter values taken from the literature. These derivations are available from the authors

Figure 6 demonstrates the influence of the agricultural tax rate, τ_A, and subsistence consumption, γ_A, on the speed of convergence to the steady state level of output per worker (the results for the speed of convergence to the steady state level of L_M are similar). When incomes are comfortably away from subsistence (i.e., $\gamma_A = 0.187$), the effect of taxes on the speed of convergence is relatively small (panel a) provided taxes fall within an empirically plausible range: $\tau_A < 0.50$.[26] At the same time, the effect of subsistence consumption on the speed of convergence is economically significant, even when there are no distortions (panel b). Countries that start with a high ratio of subsistence consumption to income approach their steady state relatively slowly (significantly less than two percent per year), whereas countries that are comfortably away from their subsistence constraints tend to approach their steady state faster (at about three percent per year).[27] What is striking, however, is the influence of agricultural taxes on the convergence rate when incomes are closer to subsistence food requirements. For relatively high values of subsistence food consumption ($\gamma_A = 0.50$ and $\gamma_A = 0.75$) relative to initial income, taxes have a dramatic impact on the convergence rate even when the tax rate is relatively modest (panel a). For low income countries, therefore, the combined effects of a higher agricultural tax and subsistence consumption significantly retards convergence, and the calibrated magnitudes are therefore broadly consistent with our econometric estimates.

5 Concluding remarks

By examining often neglected but pervasive national economic policies on agriculture that directly affect the single largest sector in poor countries, we provide new insights on the link between sectoral distortions and economic growth. Our findings suggest that, contrary to any intended effects, policies that discriminate against agriculture reduce the very surpluses they wish to mobilize and in fact significantly retard economic growth by

upon request.

[26] In the limiting case, when $\gamma_A = 0$, the convergence properties of the model are identical to those of the Ramsey model.

[27] In interpreting these results, readers should also note that, in the absence of taxes, when $\gamma_A = 1.9$, the model economy concentrates all its resources in the agricultural sector just to meet its basic subsistence food consumption needs. For similar results in the context of a one-sector endogenous growth model, see Steger (2000).

slowing the pace of structural change.

While our findings make a strong case against taxing agriculture, they should not automatically be interpreted as an endorsement of a bias in favor of agriculture. All we can conclude from the available evidence is that, compared to taxing countries, those countries that did not discriminate against their agricultural sector had relatively better economic outcomes. It is an open question whether subsidizing countries systematically combined their pricing policies with an industrial policy that was supportive of innovation and productivity growth in agriculture. To understand these important issues, we think that future work must carefully link differences in concrete economic policies to differences across countries *and* across sectors in total factor productivity growth rates.

Appendix

A Data sources and variables

Table A.1 summarizes our data sources. Below we provide the details.

A.1 Nominal rate of protection

The most common vehicle for agricultural taxes, particularly in developing countries, is state monopolies ("marketing boards"), which purchase agricultural products from domestic producers and sell them directly in domestic and international markets. (For an extensive discussion, see Krueger, Schiff, and Valdés, 1992.) These state monopolies often add very little or no value to the final agricultural product beyond transportation and storage services. In the case of subsidies, agricultural producers receive a direct payment from the state, providing the producer with returns above the international market price. Because of their widespread use in both developing and industrial countries, direct taxes and subsidies have received considerable attention in agricultural policy debates.

A.1.1 Measures and definitions

The agricultural taxation or subsidy measure we use in this study is the *nominal rate of protection* (NRP), which is defined as the percentage by which the local producer price exceeds (or falls below) the border price. This is a very common measure of direct policy interventions; see, e.g., Anderson and Hayami (1986) and Krueger et al. (1992). In the case of agricultural taxes, the NRP is negative and domestic producers receive less than the international price per unit of output. In the case of subsidies, the NRP is positive and domestic producers receive more than the international price per unit of output. (In some cases, the original data were presented in terms of nominal coefficient of protection (NCP) defined as the ratio of the domestic producer price to the international price, and we converted these using NRP = (NCP$-$1) $\times$ 100.)

We focus exclusively on the NRP because it is closer to our conceptual framework, considerably more transparent, and can be computed for a larger group of countries. Indirect protection measures – which adjust for the effect of exchange rate overvaluation and industrial protection on agriculture's relative price – were only available for a few countries.

A.1.2 Coverage and data sources

We use a basket of exportable commodities for each country, with weights applied to each commodity in the basket whenever possible.[28] Broader coverage of commodities (such as grains and "all commodities") was limited in country coverage, and this method provides the largest country and period coverage given available data. When available, we found a close association between the NRP for all commodities and the NRP for exportables only (see also Schiff and Valdes 1992).

Final data on nominal rates of protection are averages within two sub-periods: 1960–72 and 1976–84 (but see below individual country coverage). When data sources overlapped, we used the following hierarchy: SV, AH, L and TA (see table A.2 for full references) because we wanted to uniformly use exportables and achieve a broader coverage of commodities – except in the case of Republic of Korea where we used AH rather than SV (which does not report any data for exportables).

Country coverage and data sources are given in table A.2. Commodity coverage and methodology used in the secondary sources are as follows:

- Anderson and Hayami [AH] (1986, table 2.5): Weighted average for twelve commodities, using production valued at border prices as weights.

- Lele [L] (1988, table 12): The estimates are based on annual NRP for exportable cash crops, but no commodity weights are available. Commodity coverage by country is as follows: Cameroon (america coffee, robusta coffee, cocoa, cotton), Kenya (coffee, tea), Malawi (dark-fired, burley, flue-cured cotton), Nigeria (cocoa palm kernel), Senegal (groundnuts, cotton), Tanzania (tobacco, cotton, coffee).

- Schiff and Valdes [SV] (1992, table 2.3): Coverage is comprehensive; see individual country studies.

[28] We recognize that not all domestically produced agricultural goods are taxed or subsidized at the same rate as the exportable cash and food crops (and we have much less to say about livestock). However, data limitations prevent us from using more comprehensive measures. Second, we think of the nominal rate of agricultural protection as a *relative* measure of distortion, and do not explicitly measure those policies that discriminate against or in favor of other major sectors.

- Tshibaka [T] (1993, table 5.6): Zaire only. The data are reported for 1971–74, 1975–79 and 1980–82. Unweighted mean of implicit rates of protection for three export crops (coffee, palm oil and cotton).
- Tyers and Anderson [TA] (1992, table 2.6 and 2.7):The estimates are the weighted averages for grains, edible livestock products and sugar, using production valued at border prices as weights.
- World Bank [WB] (1981, p. 56): The estimates are based on NCP for selected exportable cash crops, but no commodity weights are available. Commodity coverage by country is as follows: Mali (cotton, groundnuts), Sudan (cotton, groundnuts, sesame), Togo (cocoa, coffee, cotton).

Table A.3 presents the NRP data for each country.

A.2 Income

Real GDP per capita in current prices: (cgdp) Data are from Heston, Summers, and Aten (2002, Penn World Tables Version 6.1). Data for (former West) Germany are scaled up using RGDP85 (see below).

Real GDP per capita in constant prices: Data in constant, chained 1996 prices (rgdpch) are from Heston, Summers, and Aten (2002, Penn World Tables [PWT] Version 6.1). Data are mostly available between 1960 and 2000, except that the data starts late for Germany (1970) and ends early for Congo (1997) and Taiwan (1998). There is no data on Sudan (except for 1996). Our final real GDP per capita in 1996 prices (RGDPCH96) fills in the missing observations in rgdpch with those from RGDPINT95 using growth rates. This method is used to compute 2001 and 2002 real GDP values for all countries, and data for Congo between 1998 and 2002. Data on Sudan between 1975 and 1995 and between 1997 and 2002 are obtained by using the implied growth rates in RGDPINT95, and data between 1960 and 1974 are back-casted using the implied growth rates in RGDP85 (see below). We refer to these series below as the "augmented" constant, chained series in 1996 prices. In the augmented series, all countries thus have data from 1960 to 2002, except Taiwan which ends in 1998. (For Germany (DFA) we have data from 1960 to 1992 in international 1985 prices; see RGDP85 above).

Real GDP per capita in constant 1985 dollars: (RGDP85) For the Federal Republic of Germany (DFA), and for Sudan we lack cgdp and rgdpch data from the Penn World Tables (see above), so we use rgdp85 data to fill in the missing observations. Data are in international prices (base year 1985) primarily from the Penn World Tables 5.6 and are obtained from the Global Development Network (GDN) Growth Database accessed at http://www.worldbank.org/research/growth/GDNdata.htm#5.

Annual real GDP per capita growth rate: We used the augmented constant, chained series in 1996 prices (RGDPCH96) to calculate period compound growth rates and standard deviation (volatility) of annual growth rates over two periods: 1960–72 and 1976–84. For Germany (DFA) we used RGDP85. Compound growth (g) is calculated by using endpoint observations ($X(0)$ and $X(t)$) and solving the equation $X(t) = e^{gt}X(0)$.

A.3 Economic structure

Investment share of GDP: Data are in current prices and percentages. Data are from the Penn World Tables 6.1 (ci). German data starts in 1970. Since PWT has no data on Sudan (except in 1996), for the period from 1970 to 1987, we used the sum of private and public investment as a share of GDP from the Global Development Network Growth Database, and for the period from 1996 to 2000, we used gross fixed capital formation as a percentage of GDP from the World Bank's World Development Indicators, WDI 2004. For Sudan we used 1987 data for 1990, and for Zaire we used 1997 for 2000.

Agriculture value added per worker: (ALP) Data are in constant 1995 US dollars and are taken from the WDI (2004). Agricultural value added is highly variable in the short run. So, when the data were available, we averaged three years of observations centered on the first year of each decade since 1970 (average of 1969–71, and so on). When the beginning observation was 1970 or later we used the average of the first three available observations. In particular, Malaysia and Italy start in 1970 so we used the average of 1970–72. Australia, Canada, France, Germany (DEU), Netherlands, Portugal, UK and US start in 1971, so we used the average of 1971–73. New Zealand starts in 1977, so the observation corresponding to 1969–71 is missing. Tanzanian data starts in 1990, so we used the average of 1990–92. WDI (2004) has no data on Sudan, Switzerland and Taiwan. All other countries have data since at least 1965. There are no separate data on the Federal Republic of Germany (DFA), so we used the data on unified Germany (DEU).

We filled in the missing data on Switzerland and Taiwan using data from Anderson and Hayami (1986, table A2.1). Specifically, their numbers imply that, in 1970, agricultural labor productivity in Switzerland was about 75 percent (21.7/28.8) that of France (a country with roughly similar agricultural policies). We applied this factor to French ALP in 1970. We then used the real ALP growth rate implied by Anderson and Hayami's numbers to calculate the ALP in Switzerland in 1980. Similarly, their numbers indicate that, in 1970, agricultural labor productivity in Taiwan was about 1.7 (4.5/2.7) times the real ALP in Korea. We then repeated the procedure used for Switzerland.

Relative agricultural productivity: (RLP) Relative agricultural productivity is agricultural productivity per worker divided by GDP per capita (APL/RGDP95). Since the vast majority of agricultural products are in principle traded, comparing agricultural productivity in constant US dollars would be appropriate. Not all components of GDP are internationally traded, however. So, GDP per capita or aggregate productivity is measured in international prices. In accordance with the APL data we used three year averages of real GDP per capita centered on the first year of the decade. When data availability did not allow us to calculate the APL centered on the first year of the decade, we made a similar adjustment to GDP data.

Data in constant 1995 US dollars (RGDPUS95), and in PPP ("constant 1995 international prices", RGDPINT95) are from the WDI 2004. RGDPUS95 are available at least since 1964 except for Canada (since 1965), Germany (DEU, 1971), Mali (1967), Tanzania (1988), and Turkey (1968). Constant international price data (RGDPINT95) are available only since 1975, except for Tanzania (which is only available since 1988). Neither RGDPUS95 nor RGDPINT95 is available for Taiwan. To calculate the relative productivity measures we shifted the base year in the original series from 1996 to 1995 and called the new series RGDPCH95. To do this we used two series: Real GDP per capita in chained 1996 prices (rgdpch) and real GDP per capita in current prices (cgdp), both from Heston et al. (2002). See table A.2 for data coverage.

To fill in the missing values for Taiwan corresponding to constant U.S. dollar estimates, we used, for 1970 and 1980, real GDP per capita (in constant U.S. dollars) in Taiwan divided by U.S. real GDP per capita (15 and 25 percent respectively), both from Anderson and Hayami (1986, table A2.1). We multiplied these ratios with the respective U.S. real GDP per capita from the WDI 2004.

To check the sensitivity of our results to combining two alternative measures of real GDP per capita, we repeated our calculations using three real GDP per capita figures; constant 1995 international prices (RLPINT95), constant 1995 US dollars (RLPUS95), and chained 1995 prices (RLPCH95). (We used constant U.S. dollar series for comparison purposes only.) In accordance with the ALP data, we used three year averages of GDP per capita centered on first year of the decade (average of 1969–71, and so on). When data availability did not allow us to calculate the APL or GDP centered on the first year of the decade, we calculated relative productivity by using the first or last three commonly available observations. The correlation coefficient between RLPINT95 and RLPCH95 is very high for all three sub-periods (about .98). Not surprisingly, the correlation between RLPINT95 and RLPUS95 and between RLPCH95 and RLPUS95 are relatively low; starting from about .28 in 1979–81 for RLPINT95–RLPUS95 and .30 in 1969–71 for RLPCH95–RLPUS95 and rising to .54 in 1999–2001 for RLPINT95 and to .60 in 1998-2000 for RLPCH95. On average RLPUS95 exceeds both RLPINT95 and RLPCH95 by an economically significant margin. Finally, we examined the relationship between APL and GDP per worker, from Heston et al. (2002). The qualitative results were identical to those obtained from GDP per capita.

TFP growth in agriculture: Our coverage and sources are summarized in table A.4.

Share of agriculture in GDP: Measured as the agriculture share in total value added (GDP) in percentage form, these data are mostly from World Bank WDI (2004). Some countries have missing observations. Taiwan has no data. Switzerland has no data except 1997–2000 (inclusive). For Taiwan and Switzerland, we used Anderson and Hayashi (1986, table A2.1). Data on Tanzania starts in 1990 and on Bangladesh in 1980. Data on Sudan are missing between 1988–95 (inclusive), so we used the 1987 data for 1990. Australia, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, New Zealand, Sweden, Portugal, UK and US data start in 1971, and we used the values for 1970 instead. Data for Canada ends in 1999 which we used for 2000. New Zealand ends in 1997, which we used for 2000. (Data in Anderson and Hayashi for Germany (DFA) are very similar to those for unified Germany (DEU) in the WDI (2004).)

Share of agriculture in employment: We used the ratio of agricultural population to non-agricultural population, both of which are from the FAO's FAOSTAT database, available at http://fao.org. These estimates are decennial. There are no data for Taiwan, so we used share of the male labor force in agriculture from Anderson and Hayami (1986, table A2.1). (WDI (2004) also has data this variable, but no data are available before 1980, and data from after 1980 are missing for many countries.)

Primary schooling in 1960 and 1970: Gross enrollment ratio for primary education in 1960 and 1970. Data are based on Barro and Lee (1993).

A.4 Other measures of distortions

Qualitative data on trade orientation: World Bank (1987, table 5.1, p. 83) classifies developing countries as strongly outward oriented, moderately outward oriented, moderately inward oriented, and strongly inward oriented developing country. The classification covers 1960–72 and 1976–84. We supplemented these data by our own judgments in the case of PR China (strongly and moderately inward oriented respectively) and Taiwan (strongly outward oriented in both periods). We also used judgment for Democratic Republic of Congo (Zaire) and Malawi (moderately inward oriented in both cases and both periods) based on Collier and Gunning (1999, p. 68) and World Bank (1981, table 32) which shows that both Malawi and Zaire had a more market based (mixed or private) procurement and distribution of agricultural inputs relative to other low and middle income sub-Saharan African countries. Industrial market economies are according to World Bank (1984, pp. 213–15). We did not have coverage for Egypt, Mali, Morocco, Portugal and Togo.

Parallel market premium: Ratio of parallel market premium to official exchange rate, average between 1960 and 1972 and 1976 and 1984. Data are from GDN Growth

Database. Note that values for industrial countries are added as zero. Data for Mali does not start until 1974. There is no data on Taiwan.

The parallel market premium is a commonly used measure of international trade distortions in cross-country growth regressions (see, e.g., Easterly, 1994, and Lee, 1993). The parallel market premium is typically interpreted as an economywide distortion, uniformly affecting all the sectors in the economy exposed to international trade. It is therefore distinct from sector specific measures such as the nominal rate of agricultural protection. In fact, we find that the parallel market premium is not highly correlated with the nominal rate of protection. Table A.5 shows that the bivariate correlations between the NRP and the parallel market premium are not very strong, ranging from −0.11 during the period 1960–1972 to −0.18 during 1976–1984. One of the reasons for this low correlation is that industrialized countries are automatically assigned a zero parallel market premium (hence the distortion measure is truncated), whereas the NRP captures both positive and negative distortions (taxation versus subsidies).

Relative price of investment: Average price level of investment (purchasing power parity divided by exchange rate basis in current prices) between 1960 and 1962 and thereafter average value between three years of observations centered on the first year of the decade (i.e., average of 1969–71, and so on). Data are from Heston et al. (2002). German data starts in 1970. There is no data for Sudan except in 1996, and data for Democratic Republic of Congo (Zaire) are missing between 1998 and 2000.

We consider relative price of investment goods because it is often viewed as a measure of taxation of investment goods which creates disincentives to accumulate capital and thus hurt future economic growth. While direct measures of taxation on domestically produced investment goods are not available for a large sample of countries, previous literature has found a strong partial negative correlation between the relative price of investment goods and future economic growth (see, e.g., Easterly, 1993, and Jones, 1994) and has attributed the cross-country variation on the relative price of investment goods to distortions—specifically, domestic taxes on machinery and equipment.

We find that the correlation between the NRP and the relative price of investment goods, although negative, is not very strong. Table A.5 shows that in our cross-country data set the correlation is relatively weak (−0.25) during the first part of our sample (1960–1972), and becomes slightly stronger (−0.33) in the second part (1976–1984).

Tariff rate: We also consider discrimination against imported investment goods. Most developing countries are net importers of capital goods, and high tariff rates on these goods would retard investment, thereby slowing economic growth and structural transformation. Moreover, high tariff rates on capital goods protect domestic manufacturing industries, and are thus an indirect way to discriminate against agriculture. Our measure for this variable is own-import weighted tariff rates on intermediate inputs and capital goods (OWTI) as discussed in Lee (1993), and is obtained from Barro and Lee (1993). Table A.5

shows that the correlation between the NRP and the import-weighted tariff rates on imported intermediate and capital goods is very low (-0.09) during the period 1960–1972, and rises to -0.35 between 1976–1984, suggesting some complementarity between the NRP and the tariff on imported investment goods.[29]

B Derivation of the structural change regressions

We begin with equation (11), which determines the sectoral allocation of labor, and measure structural change by computing the changes in the sectoral composition of employment over time:

$$(1+p_t)\hat{L}_{Mt} = \frac{p_t s_{Mt}}{L_{Mt}}\,\hat{s}_{Mt} - \left(\frac{1-L_{Mt}}{L_{Mt}}\right)\frac{s_{At}}{1-L_{Mt}}\,\hat{s}_{At} + (\nu-1)p_t g_{zt}. \tag{B.1}$$

Equation (B.1) relates structural change to distortions through capital accumulation, relative productivity, and subsistence consumption effects. We link the individual terms in equation (B.1) to their empirical counterparts as follows:

$$\begin{aligned}
S_M \equiv \frac{s_M}{L_M} &= \text{ratio of investment to non-agricultural output,}\\
\hat{s}_M &= \text{growth rate of ``investment to non-agricultural output ratio'',}\\
l_M \equiv \frac{1-L_M}{L_M} &= \text{ratio of agricultural to non-agricultural employment,}\\
S_A \equiv \frac{s_A}{1-L_M} &= \text{ratio of subsistence consumption to agricultural output,}\\
\hat{s}_A &= \text{growth rate of ``subsistence consumption to agricultural output ratio'',}\\
p &= \text{ratio of non-agricultural to agricultural productivity.}
\end{aligned}$$

Here we link the relative price variable p to the ratio of productivity levels (i.e., the relative non-agricultural productivity level); see equation (10). This variable is central to our analysis, because it allows us to gauge the influence of distortions on structural change through their indirect impact on relative prices. We refer to our measure of relative prices adjusted for agricultural taxation as "with taxes", and unadjusted actual prices as "without taxes." Because the prices that producers receive can be significantly distorted by agricultural taxation, it is important to address this price distortion effect in accounting for the impact of this taxation. We therefore compute relative prices received by producers using the ratio of "after tax" value added per worker in agriculture divided

[29] The tariff rates are from Lee (1993). Unfortunately, these data only correspond to various years (unspecified by Lee) in the 1980s, so do not overlap significantly with our data. In any case, assuming (unrealistically) significant persistence in tariff rates, we computed the bivariate correlations between the NRP and the tariff rate.

by GDP per worker. The tax rate we use for agricultural value added per worker is derived from the NRP.

Also, we label $\hat{s}_A$ and $\hat{s}_M$ as growth rates, but for an exact correspondence the denominators must be multiplied by the respective employment shares. Finally, in our data, GDP in current international prices (cgdp) is measured using a consumption-based price index, so the empirical counterpart of $p \cdot S_M$, which corresponds to the first ratio on the right-hand side of equation (B.1), is the ratio of investment to non-agricultural GDP:

$$p \cdot S_M = \frac{\text{investment}}{\text{cgdp} \times \text{share of non-agricultural output in GDP}}.$$

The ratio of subsistence consumption to agricultural output is difficult to measure. We interpret the influence of this variable on labor reallocation as a parameter that potentially varies across time periods, and thus estimate the following regression equation:

$$(1 + p_{jt_0})\hat{L}_{Mjt} = b_{0t} + b_{1t} p_{jt_0} S_{Mjt_0} \hat{s}_{Mjt} + b_{2t} l_{Mjt_0} + b_{3t} p_{jt_0} + \varepsilon_{jt}, \tag{B.2}$$

where $j = 1, \ldots, J$ indexes the countries in our sample, b_{it} (where $i = 0, \ldots, 3$) are parameter coefficients that are allowed to vary across time periods, and ε_{jt} is an error term. Each time period t corresponds to a decade within which we compute the growth rates, and t_0 corresponds to the initial observation of time period t. In our sample, $t_0 = 1970, 1980$.[30] We estimate the model using OLS and control for cross-sectional heteroscedasticity.

References

Akiyama, Takamasa, John Baffes, Donald F. Larson, and Panos Varangis. 2003. “Commodity market reform in Africa: some recent experience.” Washington, D.C.: World Bank Policy Research Working Paper no. 2995.

Anderson, Kym. 2006. “Reducing distortions to agricultural incentives: progress, pitfalls, and prospects.” *American Journal of Agricultural Econics* 88: 1135-1146.

Anderson, Kym, and Yujiro Hayami, with associates. 1986. *The Political Economy of Agricultural Protection: East Asia in International Perspective.* Sydney and London: Allen & Unwin.

Barro, Robert J., and Jong-Wha Lee. 1993. “International comparisons of educational attainment.” *Journal of Monetary Economics* 32: 363–394. [Accessed at http://www.nber.org/pub/barro.lee/.]

[30] We first estimated this equation using $\hat{p}$ as an interaction term for p_{jt}, but preliminary specification tests were considerably more favorable to regression model (B.2). Unfortunately, theoretically more appropriate relative TFP growth rates (that is, agriculture versus non-agriculture) are only available for a small set of countries and only in most cases for agriculture and manufacturing; see, e.g., Martin and Mitra (2001).

Barro, Robert J., and Xavier Sala-i-Martin. 1995. *Economic Growth.* McGraw Hill.

Bates, Robert H.. 1981. *Markets and States in Tropical Africa: The Political Basis of Agricultural Policies.* Berkeley: University of California Press.

Canova, Fabio. 2004. "Testing for convergence clubs in income per capita: a predictive density approach." *International Economic Review* 45: 49–77.

Chenery, Hollis, and Moshe Syrquin. 1975. *Patterns of Development, 1950–1970.* London: Oxford University Press.

Clark, Colin, and Margaret Haswell. 1964. *The Economics of Subsistence Agriculture.* St. Martin's Press.

Collier, Paul, and Jan W. Gunning. 1999. "Explaining African economic performance." *Journal of Economic Literature* 37 (March): 64–111

Dennis, Benjamin N., and Talan B. İşcan. 2007. "Productivity growth and agricultural out-migration in the United States." *Structural Change and Economic Dynamics* 18: 52–74.

Durlauf, Steven N., and Paul Johnson. 1995. "Multiple regimes and cross country growth behavior." *Journal of Applied Econometrics* 10: 365–384.

Easterly, William. 1993. "How much do distortions affect growth?" *Journal of Monetary Economics* 32: 187–212.

Easterly, William. 1994. "Economic stagnation, fixed factors, and policy thresholds." *Journal of Monetary Economics* 33: 525–557.

Echevarria, Cristina. 1997. "Changes in sectoral composition associated with economic growth." *International Economic Review* 38: 431–452.

Fogel, Robert W. 2004. *The Escape from Hunger and Premature Death, 1700-2100: Europe, America and the Third World*. Cambridge: Cambridge University Press.

Fulginiti, Lilyan E., and Richard K. Perrin. 1993. "Prices and productivity in agriculture." *Review of Economics and Statistics* 75: 471–82.

Fulginiti, Lilyan E., and Richard K. Perrin. 1999. "Have price policies damaged LDC agricultural productivity?" *Contemporary Economic Policy* 17: 469–75.

Fulginiti, Lilyan E., Richard K. Perrin, and Bingxin Yu. 2004. "Institutions and agricultural productivity in Sub-Saharan Africa." *Agricultural Economics* 31: 169-180.

Hamilton, James D. 2001. "A parametric approach to flexible nonlinear inference." *Econometrica* 69: 537–74.

Hansen, Bruce E. 1996. "Inference when the nuisance parameter is not identified under the null hypothesis." *Econometrica* 64: 413–430.

Hansen, Bruce E. 2000. "Sample splitting and threshold estimation." *Econometrica* 68: 575–603.

Heston, Alan, Robert Summers, and Bettina Aten. 2002. Penn World Table Version 6.1, Center for International Comparisons at the University of Pennsylvania (CICUP), October 2002.

Jones, Charles I. 1994. "Economic growth and the relative price of capital." *Journal of Monetary Economics* 34: 359–382.

Kongsamut, Piyabha, Sergio Rebelo, and Danyang Xie. 2001. "Beyond balanced growth." *Review of Economic Studies* 68: 869–882.

Krueger, Anne O., Maurice Schiff, and Alberto Valdés (eds.). 1992. *Political Economy of Agricultural Pricing Policy.* 5 volumes. Baltimore: Johns Hopkins University Press.

Kuznets, Simon. 1966. *Modern Economic Growth: Rate, Structure, and Spread.* New Haven: Yale University Press.

Laitner, John. 2000. "Structural change and economic growth." *Review of Economic Studies* 67: 545–561.

Lee, Jong-Wha. 1993. "International trade, distortions, and long-run economic growth." *IMF Staff Papers* 40: 299–328.

Lele, Uma. 1988. "Agricultural growth, domestic policies, the external environment and assistance to Africa: Lessons of a quarter century," in Colleen Roberts, ed. *Trade, Aid, and Policy Reform.* Washington, D.C.: World Bank, pp. 119–198.

Lewis, W. Arthur. 1954. "Economic development with unlimited supplies of labor." *Manchester School* 22: 139–191.

Lin, Justin Yifu. 1990. "Collectivization and China's agricultural crisis in 1959-1961," *Journal of Political Economy* 98 (December): 1228-1250.

Lindert, Peter H. 1991. "Historical patterns of agricultural policy," in C. Peter Timmer, (ed.) *Agriculture and the State: Growth, Employment, and Poverty in Developing Countries.* Ithaca, London: Cornell University Press.

Martin, Will, and Devashish Mitra. 2001. "Productivity growth and convergence in agriculture versus manufacturing." *Economic Development and Cultural Change.* 49: 403–422.

Matsuyama, Kiminori. 1992. "Agricultural productivity, comparative advantage, and economic growth." *Journal of Economic Theory* 58: 317–334.

Mundlak, Yair. 2000. *Agriculture and Economic Growth.* Cambridge, Mass.: Harvard University Press.

Mwabu, Germano, and Erik Thorbecke. 2004. "Rural development, growth and poverty in Africa." *Journal of African Economies* 13 (Supplement 1): 16–65.

Ndulu, Benno J., and Stephen A. O'Connell. 1999. "Governance and growth in Sub-Saharan Africa." *Journal of Economic Perspectives* 13 (Summer): 41–66.

Ngai, Rachel L., and Christopher A. Pissarides. 2007. "Structural change in a multi-sector model of growth." *American Economic Review*, in press.

Oyejide, T. Ademola. 1993. "Effects of trade and macroeconomic policies on African agriculture," in Romeo M. Bautista, and Alberto Valdés, (eds.) *The Bias Against Agriculture.* San Francisco: International Center for Economic Growth and International Food Policy Research Institute, pp. 241–262.

Pursell, Garry, and Ashok Gulati. 1993. "Liberalizing Indian agriculture: An agenda for reform." Washington, D.C.: World Bank Policy Research Working Paper no. 1172.

Quah, Danny T. 1996. "Empirics for economic growth and convergence." *European Economic Review* 40: 1353-1375.

Ratsø, Jørn, and Ragnar Torvik. 2003. "Interactions between agriculture and industry: theoretical analysis of the consequences of discriminating agriculture in Sub-Saharan Africa." *Review of Development Economics* 7: 138–151.

Restuccia, Diego, Dennis Tao Yang, and Xiaodong Zhu. 2006. "Agriculture and aggregate productivity: a quantitative cross-country analysis." University of Toronto, mimeo.

Schiff, Maurice, and Alberto Valdés. 1992. *Political Economy of Agricultural Pricing Policy. vol. 4: A Synthesis of the Economics in Developing Countries.* Baltimore: Johns Hopkins University Press.

Schiff, Maurice, and Alberto Valdés. 2002. "Agriculture and the macroeconomy, with emphasis on developing countries," in Bruce L. Gardner and G.C. Rausser (eds.) *Handbook of Agricultural Economics, Volume 2A*, Amsterdam: North Holland.

Steger, Thomas M. 2000. "Economic growth with subsistence consumption." *Journal of Development Economics* 62: 343–361.

Takeuchi, Fumihide, and Takehiko Hagino. 1998. "Intersectoral resource allocation and its impact on economic development in the Philippines." Washington, D.C.: World Bank Policy Research Working Paper no. 1883.

Timmer, C. Peter. 1988. "The agricultural transformation," in Hollis Chenery and T.N. Srinivasan (eds.) *Handbook of Development Economics, Volume 1*, Amsterdam: North Holland.

Timmer, C. Peter. 2005. "Agriculture and pro-poor growth: An Asian perspective," Washington, D.C.: Center for Global Development Working Paper 63.

Tomich, Thomas P., Peter Kilby, and Bruce F. Johnston. 1995. *Transforming Agrarian Economies: Opportunities Seized, Opportunities Missed.* Ithaca and London: Cornell University Press.

Tshibaka, Tshikala B. 1993. "Agricultural pricing and the exchange rate in Zaire," in Romeo M. Bautista, and Alberto Valdés, (eds.) *The Bias Against Agriculture.* San Francisco: International Center for Economic Growth and International Food Policy Research Institute, pp. 93–110.

Tyers, Rod, and Kym Anderson. 1992. *Disarray in World Food Markets.* Cambridge, UK: Cambridge University Press.

Vuong, Quang H. 1989. "Likelihood ratio tests for model selection and non-nested hypothesis." *Econometrica* 57: 307–333.

World Bank. 1981. *Accelerated Development in Sub-Saharan Africa: An Agenda for Action.* Washington, D.C.: World Bank.

World Bank. 1984. *World Development Report 1984.* New York: Oxford University Press.

World Bank. 1987. *World Development Report 1987.* New York: Oxford University Press.

World Bank. 2004. *World Development Indicators 2004.* CD-ROM. Washington, D.C.

Table 1: Descriptive statistics

Variable (mnemonic)	Mean	Std. Dev.	Min.	Max.	N
a) Nominal rate of protection, %					
1960–1972 (nrp6072)	−5.209	34.415	−80.300	77.666	47
1976–1984 (nrp7684)	−4.390	40.406	−79.600	111.000	47
b) Income and economic structure					
Ln GDP per capita					
in 1970 (cgdp70)	6.892	1.075	5.229	8.677	47
in 1980 (cgdp80)	7.732	1.114	5.988	9.406	47
Average GDP per capita growth					
1960–1972 (rgdpch6072)	0.028	0.021	−0.016	0.089	47
1976–1984 (rgdpch7684)	0.015	0.023	−0.034	0.065	47
Investment share in GDP, %					
in 1970 (ci70)	19.267	9.609	4.020	38.490	47
in 1980 (ci80)	19.296	7.850	5.916	34.618	47
Relative labor productivity					
in 1970 (rlp70)	0.609	0.392	0.143	1.850	44
in 1980 (rlp80)	0.606	0.387	0.145	1.532	45
Share of agr. value added, %					
in 1970 (agrva70)	22.777	15.471	2.918	66.023	45
in 1980 (agrva80)	18.961	13.327	2.209	57.920	46
Share of agr. employment, %					
in 1970 (agremp70)	48.954	29.102	2.830	92.634	47
in 1980 (agremp80)	43.131	27.920	2.614	88.969	47
Enrolment rate in primary school					
in 1970 (p70)	0.817	0.236	0.220	1.000	46
in 1980 (p80)	0.904	0.185	0.250	1.000	45
c) Other measures of distortions					
Relative price of investment					
1960–1972 (pi6072)	70.712	42.078	28.745	296.387	46
1976–1984 (pi7684)	92.229	44.795	35.199	310.956	46
Parallel market premium					
1960–1972 (premium6072)	63.799	259.494	−0.305	1,751.613	45
1976–1984 (premium7684)	40.476	135.552	−2.073	904.435	46
Tariff rates on capital goods					
1980s (owti)	0.194	0.216	0.012	1.319	42

Notes: N is the number of observations. See appendix A for data sources.

Table 2: Convergence in GDP per capita regression results

Dependent variable: Growth of real GDP per capita, 1960–72 and 1976–84

Independent variable	Regression 1	2	3	4
a) *Estimates with no threshold effects*				
Constant	0.0117	0.0723	0.0458	0.0859
	(0.0439)	(0.0471)	(0.0462)	(0.0438)
Ln(Inv)	0.0147	0.0104	0.0100	0.0061
	(0.0044)	(0.0043)	(0.0043)	(0.0043)
Ln(primary)	0.0174	0.0144	0.0101	0.0142
	(0.0059)	(0.0058)	(0.0074)	(0.0059)
Ln(pop+0.05)	−0.0216	−0.0056	−0.0123	−0.0035
	(0.0213)	(0.0217)	(0.0216)	(0.0209)
Ln(rgdp)	−0.0097	−0.0105	−0.0095	−0.0089
	(0.0034)	(0.0033)	(0.0035)	(0.0033)
NRP		0.0211	0.0177	0.0209
		(0.0079)	(0.0078)	(0.0074)
premium			−0.0010	
			(0.0007)	
pi				−0.0117
				(0.0049)
$\bar{R}^2$	0.2177	0.2769	0.2327	0.3191
Het. p-value	0.2048	0.2433	0.3005	0.3430
b) *Threshold estimate using*				
NRP	−3.50	19.00	17.70	17.70
p-value	0.0660	0.1650	0.2240	0.4120

Notes: Panel a) reports the OLS coefficient estimates of the pooled data for the growth regression without threshold effects. Heteroskedasticity corrected standard errors of coefficients are in parentheses. “Inv” is the ratio of investment to GDP, “primary” is the gross enrollment rate for the primary education, and “rgdp” is real GDP per capita all corresponding to the initial period values. “pop” is the population growth rate, “NRP” is the nominal rate of protection, “premium” is the parallel market premium, and “pi” is the relative price of investment, all period averages. “Het. p-value” is for the null test of no heteroskedasticity. Panel b) reports OLS estimates of the threshold values. p-values after the threshold estimates are for the null test of no threshold against the alternative of threshold, allow for heteroskedastic errors (White corrected) and are based on 1,000 bootstrap replications. For ease of exposition, we multiplied by 100 the coefficient estimates on NRP, premium and pi, as well as their standard errors.

Table 3: Summary results for convergence in L_M regressions

Dependent variable: Growth of L_M, 1960–72 and 1976–84

	Regression			
Independent variable	1	2	3	4
a) *Estimates with no threshold effects*				
NRP		0.0081	0.0075	0.0081
		(0.0031)	(0.0032)	(0.0031)
premium×100			0.0044	
			(0.0243)	
pi				0.0023
				(0.0050)
$\bar{R}^2$	0.1127	0.1601	0.1428	0.1711
Het. p-value	0.5256	0.6162	0.6102	0.0015
b) *Threshold estimate using*				
NRP	−8.10	−5.90	−5.90	−5.90
p-value	0.0000	0.0000	0.0020	0.0060

Notes: Panel a) only reports the OLS coefficient estimates and (in parentheses) their heteroskedasticity consistent standard errors on the distortions variables for the convergence in L_M regressions without threshold effects. Other independent variables include a constant, Ln(Inv), Ln(primary), Ln(pop+0.05), Ln(rgdp), and the initial value of L_M. See also notes to table 2. "Het. p-value" is for the null test of no heteroskedasticity. Panel b) reports OLS estimates of the threshold values. p-values after the threshold estimates are for the null test of no threshold against the alternative of threshold, allow for heteroskedastic errors (White corrected) and are based on 1,000 bootstrap replications.

Table 4: Determinants of Structural Change

Dependent variable: Rate of structural change, $(1+p)\hat{L}_M$

a) *Pooled estimates* ($N = 86$)

	Without taxes		With taxes		
	[1]	[2]	[3]	[4]	[5]
Investment rate,	0.7097	0.7034	1.3148	1.2183	1.3066
$p\, S_M \hat{s}_M$	(0.5292)	(0.5323)	(0.9086)	(0.8731)	(0.8971)
Employment shares,	0.0161	0.0153	0.0287	0.0279	0.0290
l_M	(0.0030)	(0.0029)	(0.0056)	(0.0056)	(0.0059)
Relative productivity,	0.0069	0.0068	0.0062	0.0042	0.0057
p	(0.0034)	(0.0034)	(0.0031)	(0.0035)	(0.0033)
Protection	–	−0.0001	–	−0.0003	–
		(0.0001)		(0.0002)	
Protection squared	–	–	–	–	0.0000
					(0.0000)
Adjusted R^2	0.579	0.576	0.672	0.677	0.669
Hamilton's statistic	7.4061	24.7376	2.0831	7.6923	5.6893
p-value	0.0130	0.0010	0.1099	0.0140	0.0210

b) *Cross-section estimates with taxes* ($N = 43$)

	1970–1980		1980–1990	
Investment rate,	1.1182	0.1695	1.2708	1.6138
$p\, S_M \hat{s}_M$	(1.6841)	(1.8321)	(0.9661)	(1.0758)
Employment shares,	0.0291	0.0268	0.0206	0.0204
l_M	(0.0078)	(0.0074)	(0.0061)	(0.0063)
Relative productivity,	0.0093	0.0067	0.0062	0.0037
p	(0.0053)	(0.0054)	(0.0056)	(0.0063)
Protection	–	−0.0005	–	−0.0004
		(0.0003)		(0.0003)
Adjusted R^2	0.712	0.716	0.581	0.590
Hamilton's statistic	0.0603	2.2009	16.8979	11.5877
p-value	0.7772	0.1029	0.0020	0.0080

Note: All equations include a constant that is not reported. Heteroskedasticity-consistent standard errors are in parentheses. Hamilton's (2001) test statistic is distributed $\mathcal{X}^2(1)$ under the null hypothesis that the true relationship is linear. Bootstrapped p-values are based on 1000 draws from the same sample. See equation (15) and appendix B for the baseline equation. N is the number of observations.

Table A.1: Variable description and data sources

Variable	Description	Source
Nominal rate of protection	Percentage by which the local producer price exceeds the border price, with weights applied to a basket of commodities.	Anderson and Hayami (1986), Lele (1988), Schiff and Valdes (1992), Tshibaka (1993), Tyers and Anderson (1992), and World Bank (1981).
GDP per capita	Real GDP per capita in current 1996 purchasing power parity based prices.	Heston, Summers, and Aten (2002).
Average growth rate of GDP per capita	Average compound growth rate of GDP per capita at constant, chained 1996 prices.	
Agricultural labor productivity	Value added per worker in agriculture, in constant 1995 U.S. dollars.	World Development Indicators (2004), and Anderson and Hayami (1986).
Relative agricultural productivity	Value added per worker in agriculture divided by GDP per capita, averaged between 1970–1972, etc.	World Development Indicators (2004), and Anderson and Hayami (1986).
Investment share of GDP	Investment share of GDP in current prices, %,	Heston et al. (2002), and World Development Indicators (2004).
Share of agriculture in GDP	Ratio of agricultural value added to total value added, %.	World Development Indicators (2004), and Anderson and Hayami (1986).
Share of agriculture in employment	Ratio of agricultural population to non-agricultural population, %.	FAO's FAOSTAT database available at http://fao.org and from Anderson and Hayami (1986).
Enrollment in primary school	Gross enrollment ratio for primary education.	Barro and Lee (1993).
Population	Population in 000's.	Heston et al. (2002), and GDN Growth Database.
Qualitative data on trade orientation	Classification of developing country trade policy orientation as strongly outward, moderately outward, moderately inward, and strongly inward oriented.	World Bank (1987), and Collier and Gunning (1999).
Parallel market premium	Ratio of parallel market premium to official exchange rate averaged over 1960–1972 and 1976–1984, %.	Global Development Network Growth Database.
Relative price of investment	Average price level of investment between 1960–1962, etc., on purchasing power parity basis.	Heston et al. (2002).
Tariff rate	Own-import weighted tariff rates on intermediate inputs and capital goods.	Barro and Lee (1993).

Table A.2: Coverage for nominal protection rates of agricultural producer prices

Country	Code	Source	Period	Country	Code	Source	Period
Argentina	ARG	SV	1960-72/1976-84	Malaysia	MYS	SV	1960-72/1976-83
Australia	AUS	AH	1960-70/1975-80	Mali	MLI	WB	1971-75/1976-80
Bangladesh	BGD	TA	1965-74/1975-83	Mexico	MEX	TA	1965-74 1975-83
Brazil	BRA	SV	1969-72/1976-83	Morocco	MAR	SV	1963-72/1976-84
Cameroon	CMR	L	1970-74/1975-84	Netherlands	NLD	AH	1960-70/1975-80
Canada	CAN	AH	1960-70/1975-80	New Zealand	NZL	AH	1960-70/1975-80
Chile	CHL	SV	1960-72/1976-83	Nigeria	NGA	L	1970-74/1975-84
China	CHN	TA	1965-74/1975-83	Pakistan	PAK	SV	1960-72/1976-86
Colombia	COL	SV	1960-72/1967-83	Philippines	PHL	SV	1960-72/1976-84
Congo, Dem. Rep.	ZAR	T	1971-74/1975-82	Portugal	PRT	SV	1960-72/1976-84
Côte d'Ivoire	CIV	SV	1960-72/1976-82	Senegal	SEN	L	1970-74/1975-84
Denmark	DNK	AH	1960-70/1975-80	Sri Lanka	LKA	SV	1960-72/1976-85
Dominican Rep.	DOM	SV	1966-72/1976-85	Sudan	SDN	WB	1971-75 1976-80
Egypt	EGY	SV	1964-72/1967-84	Sweden	SWE	AH	1960-70/1975-80
France	FRA	AH	1960-70/1975-80	Switzerland	CHE	AH	1960-70/1975-80
Germany, FR	DFA	AH	1960-70/1975-80	Taiwan, China	TWN	AH	1960-70/1975-80
Ghana	GHA	SV	1958-72/1976-84	Tanzania	TZA	L	1970-74/1975-84
India	IND	TA	1965-74/1975-83	Thailand	THA	SV	1962-72/1976-84
Indonesia	IDN	TA	1965-74/1975-83	Togo	TGO	WB	1971-75 1976-80
Italy	ITA	AH	1960-70/1975-80	Turkey	TUR	SV	1961-72/1976-83
Japan	JPN	AH	1960-70/1975-80	United Kingdom	GBR	AH	1960-70/1975-80
Kenya	KEN	L	1970-74/1975-84	United States	USA	AH	1960-70/1975-80
Korea, Rep.	KOR	AH	1960-70/1975-80	Zambia	ZMB	SV	1966-72/1976-84
Malawi	MWI	L	1970-74/1975-84				

Notes: Commodity coverage and methodologies vary across countries and sources as explained in the text. Germany refers to former Western Germany.
SOURCES: **AH** = Anderson and Hayami, with associates (1986, table 2.5). **L** = Lele (1988, table 12). **SV** = Schiff and Valdes (1992, table 2.3). **T** = Tshibaka (1993, table 5.6). **TA** = Tyers and Anderson (1992, table 2.7). **WB** = World Bank (1981, p. 56).

Table A.3: Nominal rate of protection

	Period			Period	
Country	1960–72	1976–84	Country	1960–72	1976–84
Argentina	−34.0	−25.0	Malaysia	−10.8	−17.4
Australia	6.3	−3.5	Mali	−44.0	−56.5
Bangladesh	9.0	−8.0	Mexico	19.0	26.0
Brazil	−26.7	−1.3	Morocco	−34.0	−7.8
Cameroon	−50.8	−55.9	Netherlands	32.3	29.5
Canada	3.7	2.0	New Zealand	−1.0	−1.0
Chile	26.7	0.0	Nigeria	−32.8	−14.6
China	5.0	−9.0	Pakistan	20.0	−31.4
Colombia	−10.4	−4.6	Philippines	−9.1	−16.3
Congo, Dem. Rep.	−24.1	−46.2	Portugal	0.2	0.3
Côte d'Ivoire	−35.7	−43.5	Senegal	−80.3	−73.4
Denmark	8.3	22.0	Sri Lanka	−22.6	−31.4
Dominican Rep.	−32.5	−25.6	Sudan	−18.0	−38.0
Egypt	−36.4	−32.1	Sweden	53.0	51.0
France	34.3	29.5	Switzerland	77.7	111.0
Germany, Fed. Rep.	51.0	41.5	Taiwan, China	−0.7	36.0
Ghana	−40.0	−79.6	Tanzania	−50.8	−55.1
India	9.0	−7.0	Thailand	−33.4	17.7
Indonesia	−21.0	36.0	Togo	−48.7	−57.7
Italy	61.7	47.5	Turkey	9.4	−5.9
Japan	61.3	80.5	United Kingdom	28.0	20.5
Kenya	−22.3	−19.8	United States	7.0	2.0
Korea, Rep.	3.3	73.5	Zambia	1.4	−8.1
Malawi	−52.6	−57.2			

Note: See table A.2 for data sources and exact coverage for periods.

Table A.4: Coverage for agricultural TFP growth

Country	Source	Period	Country	Source	Period
Argentina		–	Malaysia		–
Australia	MM	1967-92	Mali	FPY	1960-99
Bangladesh		–	Mexico		–
Brazil		–	Morocco	MM	1967-92
Cameroon	FPY	1960-99	Netherlands	MM	1967-92
Canada	MM	1967-92	New Zealand	MM	1967-92
Chile	MM	1967-92	Nigeria	FPY	1960-99
China	L	1960-83	Pakistan	MM	1967-92
Colombia	MM	1967-92	Philippines	MM	1967-92
Congo, Dem. Rep.	FPY	1960-99	Portugal		–
Côte d'Ivoire	FPY	1960-99	Senegal	FPY	1960-99
Denmark	MM	1967-92	Sri Lanka	MM	1967-92
Dominican Rep.	MM	1967-92	Sudan	FPY	1960-99
Egypt	MM	1967-92	Sweden	MM	1967-92
France	MM	1967-92	Switzerland		–
Germany, FR		–	Taiwan, China	MM	1967-92
Ghana		–	Tanzania	FPY	1960-99
India	MM	1967-92	Thailand		–
Indonesia	MM	1967-92	Togo	FPY	1960-99
Italy	MM	1967-92	Turkey	MM	1967-92
Japan	MM	1967-92	United Kingdom	MM	1967-92
Kenya	MM	1967-92	United States	MM	1967-92
Korea, Rep.	MM	1967-92	Zambia	FPY	1960-99
Malawi	FPY	1960-99			

Notes: Methodologies vary slightly across studies.
Sources: **FPY** = Fulginiti, Perrin, and Yu (2004). **L** = Lin (1990, table 4). **MM** = Martin and Mitra (2001, table 1 and translog production function with constant returns to scale imposed).

Table A.5: Correlations between NRP and other measures of distortions

	1960–1972	1976–1984
Relative price of investment	−0.254	−0.337
Parallel market premium	−0.118	−0.185
Tariff rate on capital goods imports	−0.091	−0.356

Notes: The relative price of investment and parallel market premium are average values between 1960 and 1972 in column 1960–1972 and between 1976 and 1984 in column 1976–1984. Tariff rate on imported intermediate and capital goods are for various years in the 1980s only. See appendix A for data sources.

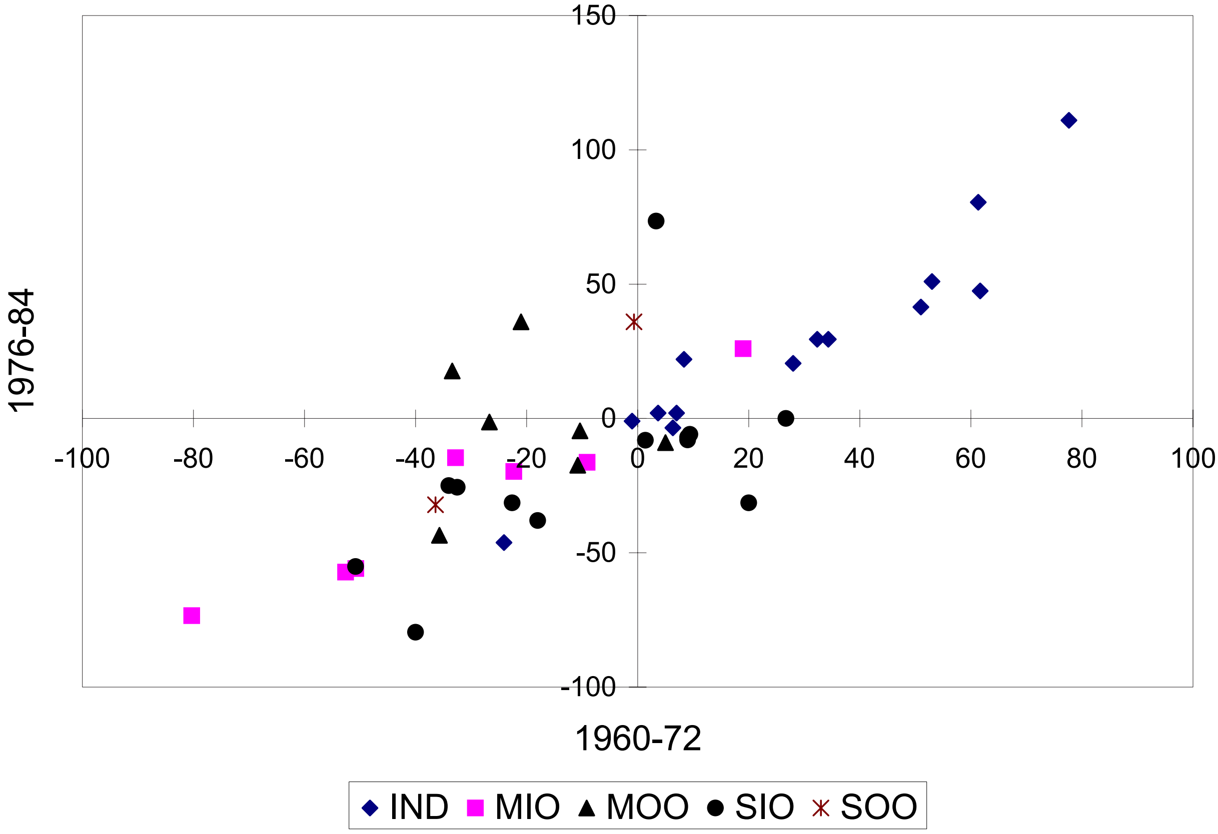


Figure 1: Nominal rate of protection

Notes: NRP is agricultural nominal rate of protection. Legend: IND = industrial market economy; MIO = moderately inward oriented developing country; MOO = moderately outward oriented developing country; SIO = strongly inward oriented developing country; SOO = strongly outward oriented developing country.
Source: For NRP see the text, and for outward orientation see World Bank (1987).

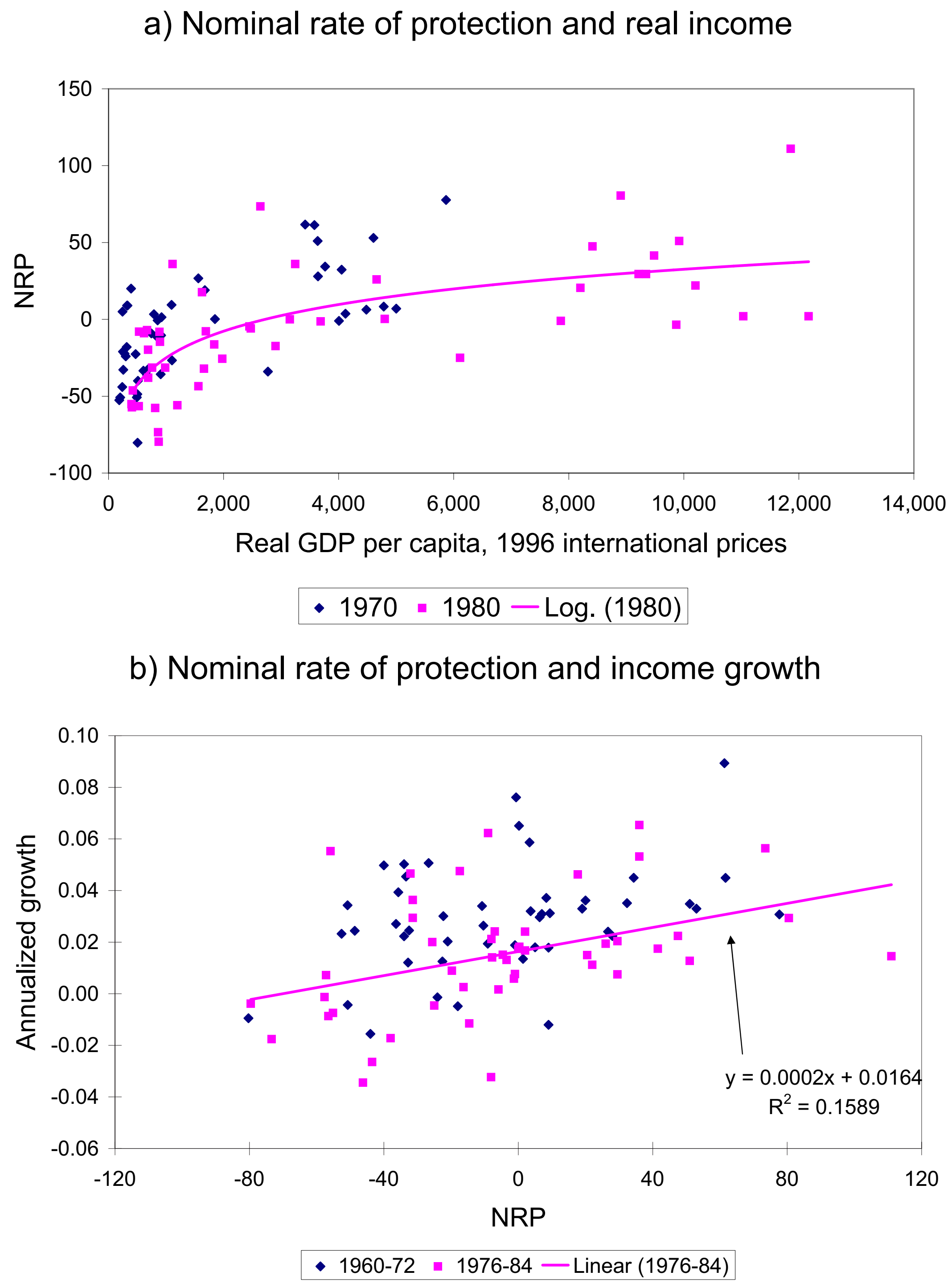


Figure 2: Nominal rate of protection and economic growth

Notes: Real GDP per capita is in constant 1996 international prices (PPP). See the text for data sources.

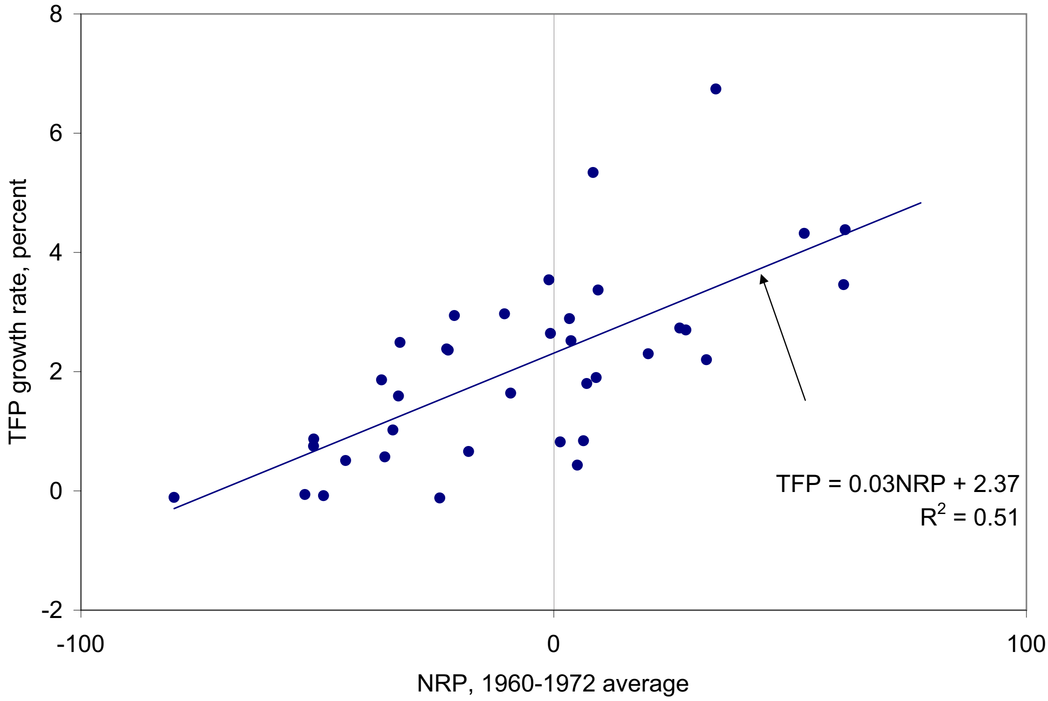
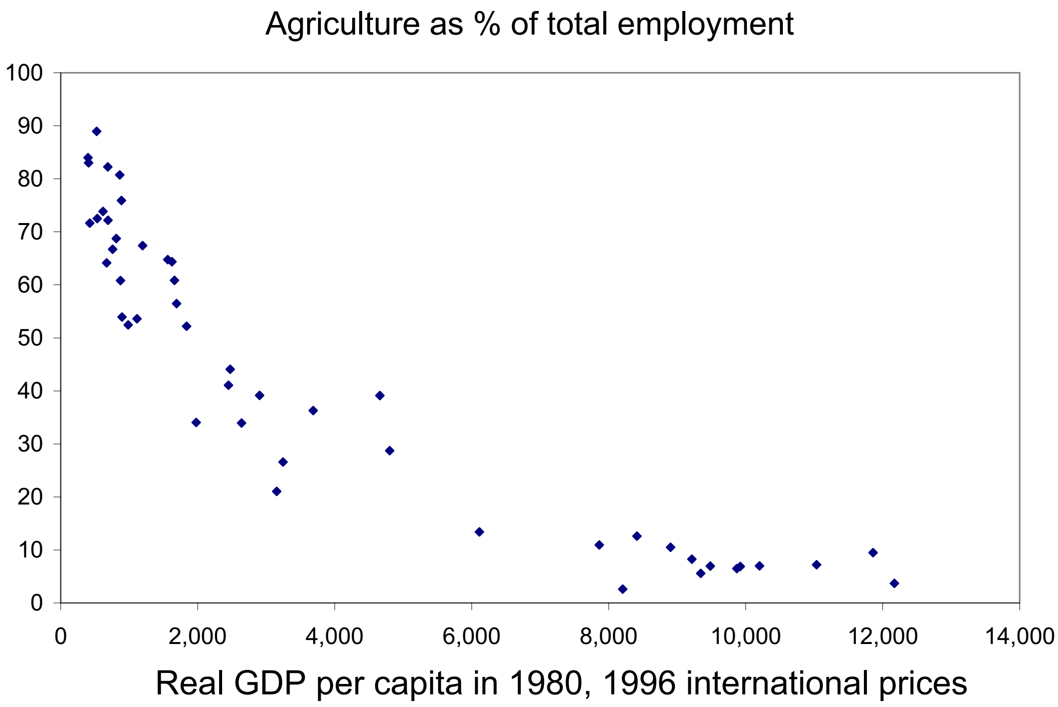


Figure 3: Nominal rate of protection and agricultural TFP growth

Notes: There are 37 countries in the sample. For coverage, see the data appendix.



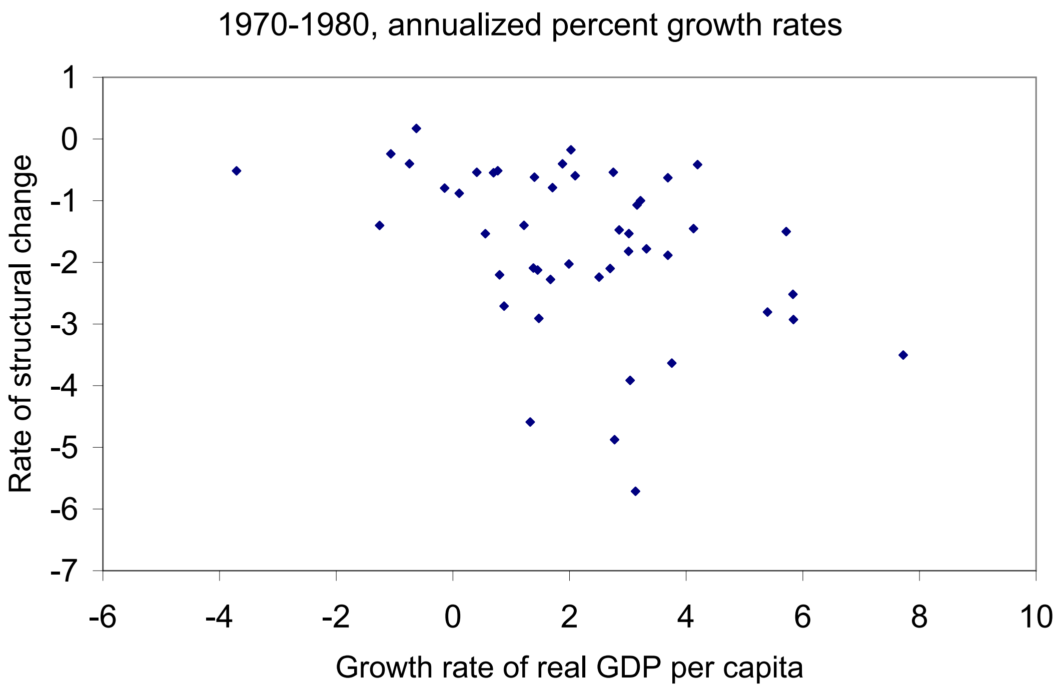


Figure 4: Structural change

Notes: The rate of structural change is measured by the annualized percent change in the share of employment in agriculture. A larger absolute value corresponds to a faster rate of reallocation of labor out of agriculture. Real GDP per capita is in constant 1996 international prices (PPP).

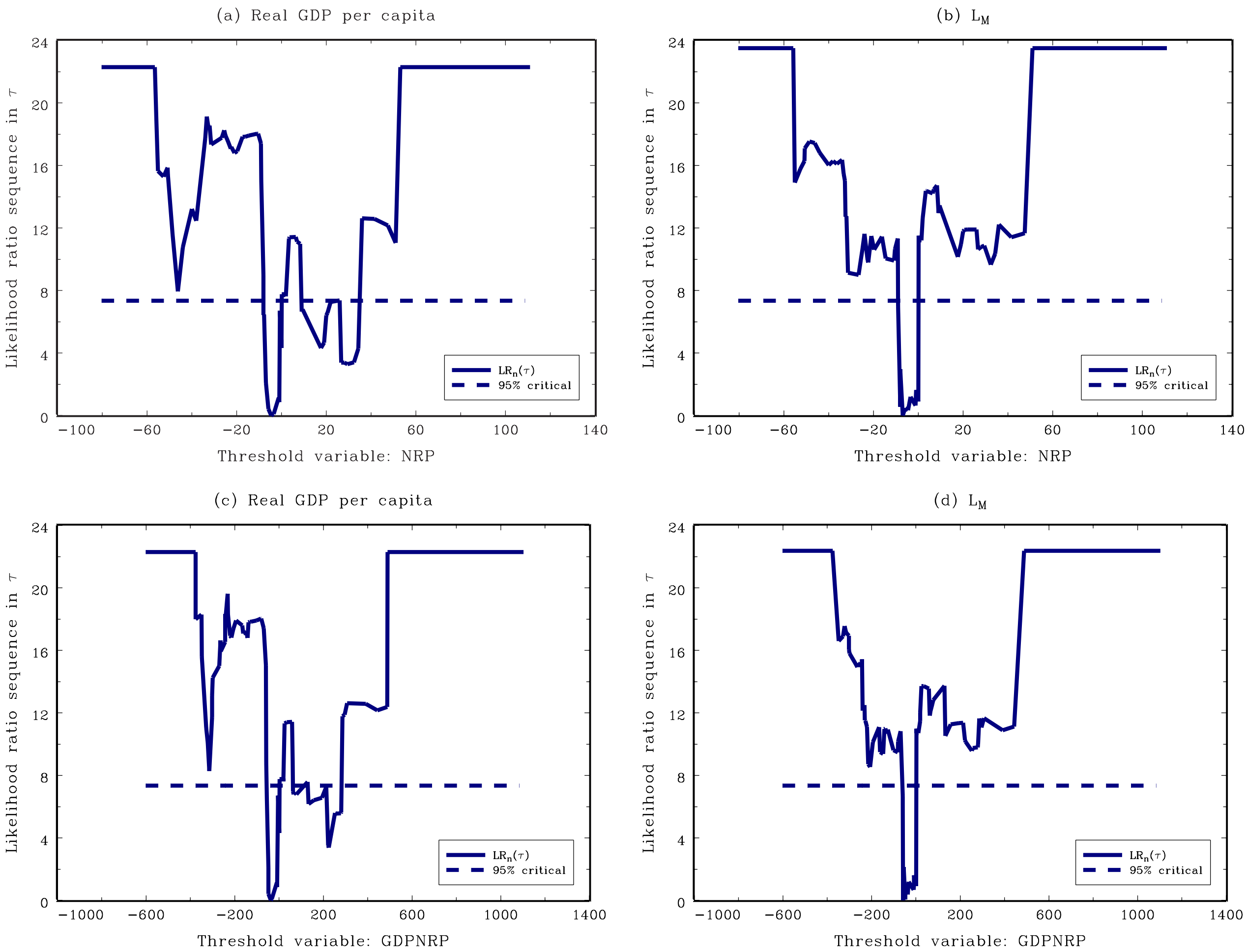


Figure 5: Confidence interval construction for threshold

Notes: "NRP" is nominal rate of protection, and "GDPNRP" is the interaction term between log GDP per capita multiplied by the nominal rate of protection. L_M is the employment share of non-agriculture.

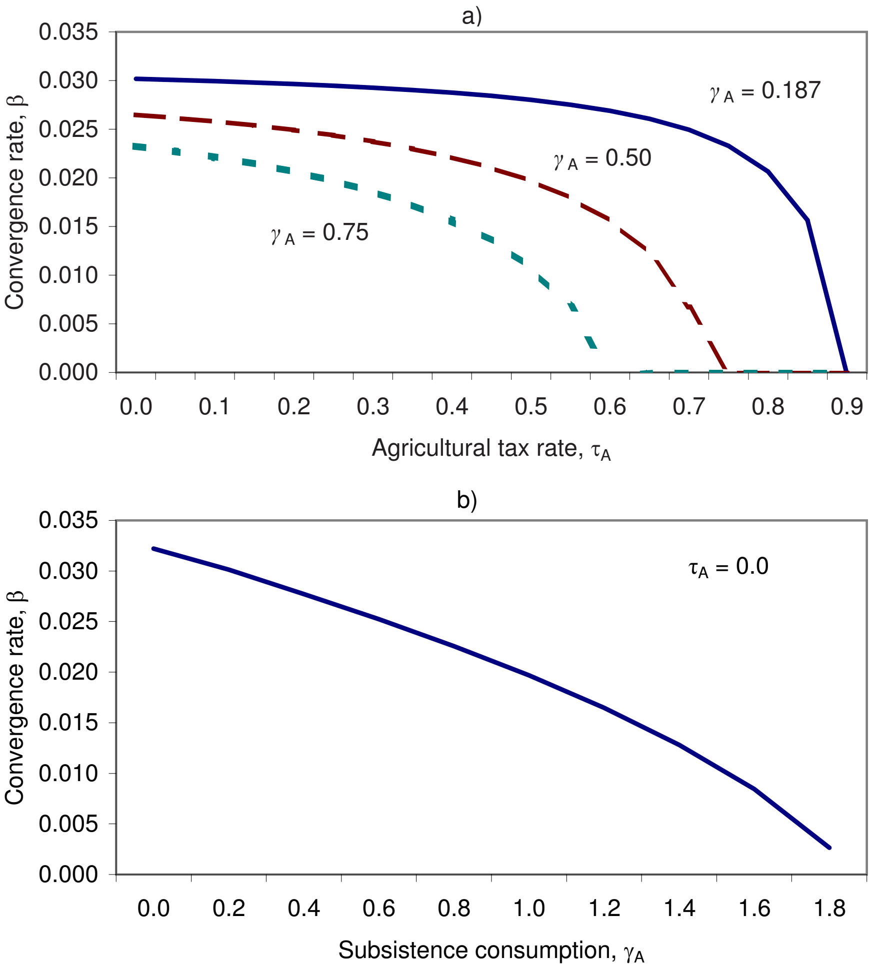


Figure 6: The speed of convergence

Notes: The parameter values are: $\alpha = 0.36$, discount rate $\rho = 0.02$, $\nu = 0.1$, $\eta = 0.82$, depreciation rate $\delta = 0.05$, $B_M = 1$, $B_A = 1$, and $Z_A = Z_M = 1$.