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Dualism and cross-country growth regressions

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Abstract

This paper develops empirical growth models suitable for dual economies, and studies the relationship between structural change and economic growth. Changes in the structure of employment will raise aggregate productivity when the marginal product of labour varies across sectors. The models in the paper incorporate this effect in a more flexible way than previous work. Estimates of the models imply sizeable marginal product differentials, and indicate that the reallocation of labour makes a significant contribution to the international variation in productivity growth.

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Keywords: Structural change, dualism, wage differentials, TFP growth.

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1 Introduction

This paper takes as its starting point three related observations, which together form a puzzle. First, the study of growth using dual economy models has a long history. Second, development economists in the 1960s and 1970s frequently discussed the role of structural change in economic growth, and especially the reallocation of labor from agriculture. Third, these twin aspects of the development process, dualism and structural change, have been almost completely absent from recent empirical growth research. Much of that research proceeds as if structural change can be ignored.[1]

With these points in mind, we investigate the implications of dualism and structural change for empirical growth models. In particular, we study dual economies in which the marginal product of labor is lower in agriculture than in the rest of the economy. This differential across sectors could arise for a number of reasons: the costs of rural-urban migration, urban disamenities, a recurring risk of unemployment in urban areas, income sharing in agriculture, or efficiency-wage considerations. It may simply be a disequilibrium phenomenon, associated with technical change or capital accumulation in the non-agricultural sector, and a less than instantaneous migration response.

If the marginal product of labor is relatively low in agriculture, moving workers to sectors where the marginal product is higher will raise total output. From the perspective of the aggregate economy, this additional output has been produced with no change in the total inputs of capital and labor. This implies that the reallocation of labor has raised aggregate productivity.[2] Our paper seeks to examine whether labor reallocation is an important source of productivity growth in practice, with an especial focus on developing countries.

We begin by setting out a simple two-sector model of a small open economy, showing that a one-sector model will be a good approximation only under restrictive assumptions. We then show how conventional growth regressions can be augmented to allow for structural change. Our empirical model allows the magnitude of the marginal product differential between agriculture and non-agriculture to vary across countries in a more flexible way than previous work.

[1] The textbooks by Bardhan and Udry (1999), Basu (1997) and Ray (1998) include discussions of dualism. Well-known studies of structural change include Chenery and Syrquin (1975) and Chenery, Robinson and Syrquin (1986). The criticism that too much growth research ignores dualism and structural change has been made by Naqvi (1996), Pack (1992), Ruttan (1998), Stern (1991) and Temple (2005) among others. Kelley and Williamson (1973) sounded a much earlier warning that conventional approaches could yield misleading findings in the context of dualism.

[2] Weil (2004, p. 284-289) provides an especially clear discussion of the aggregate effects of labor misallocation.

When we turn to the data, we find that structural change terms substantially raise the explanatory power of growth regressions. For example, we find that regressions including only structural change terms, initial TFP and regional dummies can explain around half the international variation in TFP growth; when the structural change terms are excluded, this proportion falls to a third. We also use our estimates to infer the size and cross-country variation of intersectoral differentials, and compare our results with the available microeconomic evidence.

The precise way we incorporate variation in sectoral differentials is new to this paper. We describe a set of assumptions under which the cross-section relationship between growth and the extent of structural change will be convex rather than linear.[3] This result may appear surprising, so we sketch the intuition here. Note that if wages are roughly equal to marginal products, the growth bonus associated with structural change is increasing in the size of the intersectoral wage differential. If we had to guess which countries have the largest wage differential, we might well guess those countries in which the observed extent of structural change is most rapid, reflecting large private gains from switching sectors. Conversely, in countries where structural change has recently slowed down, such as the countries of Western Europe, we might infer that wage differentials have been virtually eliminated. But this implies that the growth impact of a given extent of structural change will be greatest in those countries experiencing more rapid structural change, because these are also the countries, at least on average, in which the intersectoral differential is greatest.

At the aggregate level, this translates into a convex relationship between growth and structural change in the international cross-section, as we describe more formally below. Our estimates of the model suggest this convex relationship may be present in the data, consistent with the idea that marginal product differentials vary systematically across countries. The estimates suggest that, for some countries, the differentials are similar in magnitude to the rural-urban wage gaps observed in microeconomic data. Although the empirical estimates are consistent with a significant extent of dualism, we also find evidence that its importance has declined over time.

Various objections to this exercise can be raised, and we will discuss many of them later. We view our cross-country approach as a complement to micro-

[3] Here and throughout the paper, we use the term "structural change" in a narrower sense than usual, to refer only to changes in the employment share of agriculture. Dowrick (1989) studies additional forms of structural change in OECD countries. He finds marginal product differentials between agriculture and non-agriculture, but cannot reject the hypothesis that the marginal product of labour in industry is equal to that in services.

economic observations on rural-urban wage gaps. Note that wage gaps are not directly informative about the extent of differentials in the (unobserved) marginal product of labor. Wages may not be equal to marginal products for a wide variety of reasons, and the microeconomic evidence is potentially misleading in other regards. If we want to investigate the possible extent of marginal product differentials, or quantify the associated effect of structural change on growth, then cross-country regressions seem worth exploring.[4]

We should emphasize that the paper does not provide a complete account of the role of structural change in allowing growth to take place. In the absence of reallocation, disequilibrium across sectors would steadily increase, and output would be lower than in the case of smooth adjustment. The present paper does not seek to assess this "permissive" role of structural change in growth, despite its obvious importance. One reason for this omission is that the broader question may not be well posed. Structural change is an endogenous process, driven by sectoral productivity growth, income elasticities of demand, and changes in factor endowments and world prices, among other forces. Given that sectoral structure is clearly a general equilibrium outcome, to ask the question "What is the growth effect of structural change?" is too much like asking "What is the growth effect of equilibrium prices and quantities?". We therefore restrict attention to a narrower and well-defined question, namely the direct contribution of labor reallocation to aggregate productivity growth in economies characterized by sizeable differentials in the marginal product of labor.

The structure of the paper is as follows. Section 2 sketches the basic ideas and relates our work to the existing literature. Section 3 introduces the empirical model. Section 4 presents some stylized facts about dualism and structural change. Sections 5 and 6 report estimates of growth regressions and TFP growth regressions, and robustness tests. Section 7 presents instrumental variable estimates based on 2SLS, GMM and Fuller's modification of the LIML estimator. Section 8 examines the pattern of marginal product differentials implicit in our empirical results. Finally, section 9 rounds off with brief conclusions.

[4]There is a possible analogy here with the empirical literature on education and growth. It is well known that studies of this relationship at the aggregate level are faced by serious problems, and that for most purposes it is better to estimate the returns to education more directly, using microeconomic data. On the other hand, it is hard to draw conclusions about the direct impact of education on productivity without estimating production relationships at some level of aggregation (whether firm, industry, region or country).

2 Previous research

Our paper is founded on the idea that the marginal product of labor may be higher in urban non-agriculture than in rural agriculture. This is a narrow intepretation of "dualism" relative to the long tradition of dual economy models, and we do not address all possible consequences of dualism for the specification of growth regressions.[5] Instead, we revisit an older literature that links structural change and growth, ideas that are noticeably absent from the burst of empirical studies that followed Barro (1991) and Mankiw, Romer and Weil (1992).

In the older literature, some of the best known contributions include Denison (1967, 1974) on the postwar growth of developed countries. Similar ideas also appeared in Kuznets (1961) and are briefly discussed in Barro (1999). These authors show how growth accounting decompositions should include structural change terms, usually the rate of change of a sectoral employment share. The main drawback of Denison's approach and its extension in Temple (2001) is that the magnitude of the intersectoral wage differential has to be imputed, based on an educated guess of one form or another.[6]

The same ideas can be used to derive specifications for cross-country growth regressions, as in a pioneering study by Robinson (1971) and influential contributions by Feder (1983, 1986). In this approach, the researcher treats the structural change term as an explanatory variable, and estimates its coefficient from the data. This removes the need for guesswork about the extent of differentials, at the expense of introducing other problems. Feder's empirical model is derived using assumptions about the relationship between the marginal products of labor in each sector and economy-wide per capita output, but it is possible to derive similar results under less restrictive assumptions, as we will demonstrate below.

Recent research on structural change and growth has focused mainly on theory, especially concerning the long-run evolution of sectoral structure. Recent papers studying multi-sector models include Acemoglu and Zilibotti (1999), Atkeson and Kehoe (2000), Bencivenga and Smith (1997), Caselli (2005), Caselli

[5] See Temple (2005) for a broader discussion. Our analysis is closer to models of "modern sector dualism" (or an imperfect labour market) than "traditional sector dualism" (where the wage exceeds the marginal product in agriculture, or the agricultural wage is independent of labour demand in the modern sector). This classification of dual economy models is due to Bertrand and Squire (1980).

[6] Related methods for quantifying the effect of resource reallocation have been derived by Lipsey and Carlaw (2004), Syrquin (1984, 1986) and Pack (1992). Syrquin's method uses data on sectoral outputs and inputs and the capital share to derive what he calls the net allocation effect. The method provides a convenient lower bound on the importance of reallocation, but the required data are not always available for developing countries.

and Coleman (2001), Chanda and Dalgaard (2005), Dennis and İşcan (2004), Doepke (2004), Echevarria (1997), Falkinger and Grossman (2005), Galor, Moav and Vollrath (2005), Galor and Mountford (2004, 2006), Galor and Weil (2000), Gemmell and Lloyd (2002), Gollin, Parente and Rogerson (2002, 2004), Graham and Temple (2006), Greenwood and Seshadri (2005), Greenwood and Uysal (2005), Gylfason and Zoega (2004), Hansen and Prescott (2002), Humphries and Knowles (1998), Jeong and Townsend (2005), Kongsamut, Rebelo and Xie (2001), Laitner (2000), Lucas (2004), Ngai and Pissarides (2004), Paci and Pigliaru (1999), Restuccia, Yang and Zhu (2006), Robertson (1999), Temple (2001, 2005), Temple and Voth (1998) and Weisdorf (2006). Some of these papers have a quantitative component; for example, Gollin, Parente and Rogerson (2002, 2004) investigate the role of agriculture and home production in long-run development, using calibrated models.

The papers closest to ours are those of Dowrick and Gemmell (1991), Landon-Lane and Robertson (2003) and Poirson (2000, 2001). As in our paper, these authors consider the implications of structural change for growth regressions, but our contribution differs in a number of respects. Above all, we use estimates of a relatively flexible model to infer not only the magnitude of marginal product of differentials, but also the extent to which differentials might vary across countries.

Our study of growth in the presence of marginal product differentials can be seen as a dynamic counterpart to the recent analyses of Temple (2004) and Vollrath (2005). Their work examines the static output losses that are associated with factor market distortions, by comparisons of a distorted equilibrium with a first-best allocation in which marginal products are equalized across sectors. Vollrath's work, in particular, suggests that distortions can have sizeable effects on aggregate productivity. Our analysis studies the dynamic implications of the same idea. Since our estimates imply that differentials in marginal products are substantial, Vollrath's emphasis on the possible importance of factor market distortions seems justified.

Also relevant, Caselli (2005), Chanda and Dalgaard (2005) and Restuccia, Yang and Zhu (2006) include discussions of aggregate TFP that take into account cross-country differences in sectoral composition. One of Caselli's calculations suggests that if all countries retained their current employment allocations, but shared the same sectoral TFP levels as the USA, international disparities in output per worker would be greatly reduced. There is still scope for misallocation of labor to explain some of the remaining disparities, however. Restuccia, Yang and Zhu analyse the role of agricultural productivity in especial detail, and their calibration exercise allows for marginal product differentials between sectors. In the absence of direct estimates of differentials of the kind we provide in this paper, they use the ratio of average products across sectors as a proxy, an approach that we discuss in section 4 below.

Our contribution is more distantly related to a long history of theoretical work on aggregation. The main aggregation result that macroeconomists are familiar with is that, if all firms use the same production technology, face the same factor prices, and use inputs efficiently, then the aggregate production function will just be a scaled-up version of the firm-level production functions. The simplicity of this 'representative firm' approach is appealing, but in a two-sector world the task of aggregation is more complicated. This is so even if we assume that capital and labor are homogeneous, and factor returns equalized across sectors. If these inputs are efficiently allocated, to maximize total output, the values of maximized output at given combinations of capital and labor will trace out a surface that can be thought of as an aggregate production function.[7] However, this function may not be simple in form. It is easy to show that if two sectors each have Cobb-Douglas production technologies, and if the exponents on inputs differ across sectors, the aggregate production function cannot be Cobb-Douglas.[8]

Since in this paper we assume a marginal product differential between sectors, aggregation is even less straightforward, because the allocation of factors across sectors is no longer efficient. The next section will reaffirm that a two-sector economy is unlikely to be well approximated by an aggregate production function, except under restrictive assumptions. Formally, the lack of such a function implies that "aggregate TFP" is not a well-defined concept. For simplicity, but with some loss in precision, we will continue to use the term "TFP growth" as a convenient shorthand for the Solow residual obtained by a growth accounting exercise based on aggregate data. As we will see in the next section, this residual has a number of components in a two-sector model, including not only a share-weighted average of the sectoral TFP growth rates, but also terms that capture the effect of structural change.

[7] The efficient allocation of factors is crucial here, as pointed out by May (1946) and Pu (1946). For general treatments of aggregation problems, see Blackorby and Schworm (1988) and Fisher (1992), or Felipe and Fisher (2003) for an accessible review. Temple (2006) discusses the role of aggregation in growth economics.

[8] The way to see this is to write down the aggregate labour share as a share-weighted average of labour shares in the two sectors. If the output shares of the sectors change, the weights and the aggregate labour share will change. Hence there cannot be an aggregate Cobb-Douglas production function, since that would imply a constant labour share at the aggregate level.

3 Deriving an empirical growth model

This section first describes a measure of the extent of structural change, and then develops an empirical model that relates aggregate productivity growth to structural change. Recall that our model will build on a simple idea: countries which exhibit rapid structural change are also likely to be the countries in which the intersectoral wage gap is relatively large. This idea leads to a framework for analysing reallocation effects that is more flexible than previous contributions. Before deriving the main results, we should also note a maintained assumption of our analysis. We will assume that any observed effects of reallocation arise because of marginal product differentials, rather than sector-specific production externalities. Where present, these externalities would be likely to lead to a relationship between growth and the extent of structural change, even in the absence of a marginal product differential.

To measure the extent of structural change, we define:

$$p = -\frac{\Delta a}{a} \tag{1}$$

where a is the share of agricultural employment in total employment. We call this the 'migration propensity', denoted p. This can be interpreted as the proportion of agricultural workers who migrate in a given period if we assume that, in the absence of migration, the labor forces in the two sectors would grow at the same rate (although this assumption is not needed for our empirical work).

Our empirical framework will assume that the propensity to migrate depends on the ratio of wages in the two sectors. In one of the models we consider, we assume that migration ceases when the intersectoral wage ratio falls to a level denoted by k, initially assumed to be the same across countries. Hence in a long-run migration equilibrium, wages in the two sectors (w_a in agriculture and w_m in non-agriculture) are related as follows:

$$w_m = kw_a \tag{2}$$

where $k \geq 1$.[9] The possibility that $k > 1$ could be motivated in various ways, and Temple (2005) presents several models in which a fixed wage differential emerges as an equilibrium outcome.

We now require an equation that relates the extent of structural change to the wage ratio. A key assumption is that the strength of this response is roughly the

[9] We assume that migration only ever takes place in one direction, towards non-agriculture. There are instances where agriculture's share of employment has increased over short horizons, especially at times of economic crisis. But over the longer time periods used in our empirical work, the agricultural employment share has fallen for all the countries we consider.

same across countries. Under this assumption, we can use the observed extent of structural change to infer the magnitude of the wage differential, and hence the growth impact of a given employment shift. To implement this empirically, we will restrict attention to models where workers base their migration decisions only on the current ratio of wages in the two sectors.[10] The particular functional form we choose is:

$$\begin{aligned} p &= \frac{x}{1+x} \qquad (3) \\ \text{where } x &= \psi\left(\frac{w_m}{kw_a} - 1\right) \end{aligned}$$

where the parameter ψ captures the speed of adjustment to the long-run equilibrium, initially assumed to be constant across countries. One possible interpretation of (3) is that it reflects job search by agricultural workers, where p is the probability of a successful match with an urban firm, and this match probability is increasing in the intensity of search, which in turn is increasing in the intersectoral wage ratio.

This simple assumption allows us to derive an empirical model that is easy to interpret and can be estimated by least squares. We start by using (3) to derive an equation for the modern sector wage in terms of the agricultural wage and p, k and ψ. Note that

$$x = \frac{p}{1-p} = \psi\left(\frac{w_m}{kw_a} - 1\right)$$

so the 'odds ratio' for migration is increasing in the wage gap between the two sectors. Rearranging we have:

$$\frac{w_m}{w_a} = k\left(1 + \frac{1}{\psi}\frac{p}{1-p}\right) \qquad (4)$$

where the second term in the brackets is zero in a long-run migration equilibrium (when $p = 0$). Hence the specification captures the intuition referred to earlier. Under the assumption that the speed of adjustment (ψ) and the equilibrium differential (k) are similar across economies, we can infer the extent of the current wage ratio (w_m/w_a) using information on the observed pace of structural change, as measured by p.

[10] This is obviously a simplification, since the migration decision is likely to be forward-looking. The role of expectations is difficult to capture in a model that can be taken to the cross-country data, however. Our simplification may be reasonable if workers are impatient or adjustment is slow, and can be justified by our later empirical finding that reductions in differentials over time are noticeable but not dramatic. Use of the current wage ratio is not uncommon in theoretical work, as in Neary (1978, p. 674) and Mas-Colell and Razin (1973, p. 75) and the references therein.

We now investigate the empirical implications of equation (4). We consider a simple model of a small open economy, essentially a general equilibrium model of production with two sectors and two factors, as in the 2 x 2 model of textbook trade theory. The two sectors are rural agriculture and an urban non-agricultural sector, both perfectly competitive. The output of both sectors can be traded on world markets, but the economy is closed to international movements of capital and labor.

The agricultural good is the numeraire. Our assumptions imply that world prices tie down the relative price of the modern sector good, and we denote this relative price by q. In real terms, total output is then given by

$$Y = \frac{Y_a + qY_m}{\Omega(1,q)} \tag{5}$$

where Y_a and Y_m are output quantities in agriculture and non-agriculture respectively, the numerator is nominal output, and the denominator is a GDP price deflator.

Output in each sector is produced by capital and labor. The production functions in the two sectors have constant returns to scale and are given by:

$$\begin{aligned} Y_a &= A_a F(K_a, L_a) \\ Y_m &= A_m G(K_m, L_m) \end{aligned} \tag{6}$$

where A_a and A_m are total factor productivity in agriculture and non-agriculture respectively. We assume that workers are paid the values of their marginal products, so we have:

$$\begin{aligned} w_a &= A_a F_L \\ w_m &= qA_m G_L \end{aligned} \tag{7}$$

where the L subscript denotes the partial derivative with respect to labor. Capital also receives its marginal product in both sectors, and any difference in rental rates is immediately eliminated, so using the same notation we have:

$$A_a F_K = qA_m G_K = r \tag{8}$$

where r is the rental rate on capital (ignoring depreciation for simplicity). We denote the aggregate labor share by η and the capital share by $1-\eta = rK/Y$.

It will also be useful to define a variable $\phi = w_a L/Y$ which is approximately equal to the labor share.

We start with the simple case where there is no wage differential ($w_m = w_a$). The necessary results are then a special case of those in Jorgenson and Griliches (1967). As usual with growth accounting decompositions, the results are easiest to derive in continuous time. Growth in aggregate output, using Divisia indices for both prices and quantities, can be written as:

$$\frac{\dot{Y}}{Y} = s(t)\frac{\dot{Y}_a}{Y_a} + (1 - s(t))\frac{\dot{Y}_m}{Y_m} \tag{9}$$

where $s(t)$ is the nominal output share for agriculture at time t, or $s(t) = Y_a/(Y_a + qY_m)$. The expression (9) shows that real output growth is a weighted average of the growth rates of real quantities, where the weights are the shares of each sector in value added.

Again using standard results, the aggregate Solow residual is given by:

$$\begin{aligned} \frac{\dot{Z}}{Z} &= \frac{\dot{Y}}{Y} - (1 - \eta(t))\frac{\dot{K}}{K} - \eta(t)\frac{\dot{L}}{L} && (10) \\ &= s(t)\frac{\dot{A}_a}{A_a} + (1 - s(t))\frac{\dot{A}_m}{A_m} && (11) \end{aligned}$$

The first equality is the conventional expression for the Solow residual: output growth minus a weighted average of input growth rates, where the weights are equal to the aggregate factor shares. This simplicity is slightly deceptive, however. Given the two-sector structure of our model, the aggregate factor shares will tend to vary across countries and over time, even if the sectoral production functions are both Cobb-Douglas. This is because the aggregate factor shares are weighted averages of the sectoral factor shares, with weights equal to the shares of each sector in total value added.

The second equality shows that the aggregate Solow residual is simply a weighted average of the sectoral TFP growth rates, where the weights are equal to the output shares. This can be seen as a special case of the more general principles of Domar aggregation (for example, Jorgenson and Stiroh 2000). Although straightforward, the result arguably deserves wider attention, because empirical growth research has often assumed that efficiency growth is the same across countries, as in the influential contribution of Mankiw, Romer and Weil (1992). Their justification is that technologies can be transferred across national borders. In a two-sector world, this argument no longer goes through, except in unlikely special cases. Aggregate efficiency growth is unlikely to be the same for all countries, even when they all have access to the same technologies.[11]

Having established these basic principles, we now turn to the case of an intersectoral wage differential. The appendix shows that our assumptions lead to (10) and the following modification of (11):

$$\frac{\dot{Z}}{Z} = s(t)\frac{\dot{A}_a}{A_a} + (1 - s(t))\frac{\dot{A}_m}{A_m} + (k-1)\,\phi(1-a)\frac{\dot{m}}{m} + k\phi\frac{1}{\psi}\frac{p}{1-p}(1-a)\frac{\dot{m}}{m} \quad (12)$$

where $m = 1 - a$ is the share of non-agricultural employment in total employment.

This provides the decomposition of the aggregate Solow residual that is at the heart of our later empirical work. When there is no wage differential ($k = 1$) and the adjustment response to disequilibrium is instantaneous ($\psi \to \infty$) then the equation collapses to the original form in (11). Otherwise, labor reallocation makes a direct contribution to growth in aggregate productivity. The last two terms of (12) represent the 'growth bonus' obtained by reallocating labor to a sector where its marginal product is higher. It generalizes the earlier results of Kuznets (1961) and Denison (1967). Since the migration propensity p is related to the extent of structural change as measured by $\dot{m}/m$, equation (12) implies a convex relationship between growth and structural change, as sketched in the introduction to the paper. Implicitly, the wage differential (w_m/w_a) is allowed to vary across countries with different values of p, in the way described by equation (4).

In practice the two structural change terms are likely to be highly correlated. Our empirical work will often use restricted models, where we drop one of the two terms and examine the effect of the other. The first option is to assume the wage differential is the same across countries and time, in which case the last term is no longer relevant and we are back at the equations of Kuznets (1961) and Denison (1967). The second option is to assume that there is no wage differential in equilibrium, so that $k = 1$ and the first structural change term vanishes. We will experiment with both specifications in the empirical work that follows, and show that the second option (implying varying differentials, but only in disequilibrium) tends to perform slightly better.

We now turn to some of the details of implementing this approach empirically. The simplest approach is to estimate equations based on (12) using aggregate TFP growth rates previously calculated by Klenow and Rodriguez-Clare

[11] In principle one could imagine a long-run equilibrium in which all countries converge to the same sectoral structure. But this, too, is likely to require some restrictive assumptions, and such a long-run outcome is unlikely to be relevant over the time spans considered in our regressions. For relevant empirical work, see Wacziarg (2001) and Imbs and Wacziarg (2003).

(1997), Bernanke and Gürkaynak (2001) and Bosworth and Collins (2003). In this case the regression specification can be thought of as:

$$\frac{\dot{Z}}{Z} = \beta' X + (k-1)\,\phi MGROWTH + k\phi\frac{1}{\psi}DISEQ \tag{13}$$

where X is a vector of determinants of aggregate TFP growth, and the structural change terms (the explanatory variables) are

$$\begin{aligned} MGROWTH &= (1-a)\frac{\dot{m}}{m} \approx \Delta m \\ DISEQ &= \frac{p}{1-p}(1-a)\frac{\dot{m}}{m} \approx \frac{p}{1-p}\Delta m \end{aligned}$$

We can therefore test whether the structural change terms explain variation in aggregate TFP growth across countries. There are some approximations involved: as we have seen, aggregate TFP growth must depend on sectoral TFP growth rates, and the value added shares, as well as the structural change terms. But since it is very difficult to measure capital stocks at the sectoral level, we effectively treat sectoral TFP as unobservable, and rely on the vector X to capture the cross-section variation in aggregate TFP growth that is not due to structural change. In our empirical work, we will use regional dummies and the initial level of aggregate TFP in this role. We sometimes also add the value added share $s(t)$ to this list, since then we might obtain unbiased estimates of the structural change parameters, if the difference between the two sectoral TFP growth rates is the same for all countries.

A second approximation is that, even when the wage ratio is allowed to differ across countries (via $DISEQ$) it must be assumed constant within each country over the time period of the regression. We later examine this assumption empirically. In principle, estimation of the equation for subperiods can reveal whether differentials have changed over time.

The regression (13) is simple and easy to implement. We also use an alternative strategy that has a less direct connection to the theory, but is potentially informative. As we have seen, models for growth in GDP per worker should allow productivity growth to vary across countries. To analyse this in more detail, we take the empirical growth model derived by Mankiw, Romer and Weil (1992) (MRW from now on) and extend it to include structural change terms. Although this approach again involves some approximations, described in the appendix, it also has a number of strengths. First, unlike measuring TFP growth by accounting methods, it does not require any capital stock data. This is a major advantage given that constructing reliable measures of the capital stock for developing countries is a difficult task (Pritchett 2000). Second, we can investigate the extent to which structural change terms raise the explanatory power of some well-known empirical growth models. It turns out that allowing for structural change raises the explanatory power of these regressions substantially. Third, we can also see whether the introduction of structural change terms modifies previous conclusions from growth regressions.

Importantly, our main findings are robust to using either output growth or TFP growth as the dependent variable. The one exception to this will arise in the instrumental variable estimates, where (surprisingly) we find it easier to obtain precise estimates of structural change effects when the dependent variable is output growth rather than TFP growth.

Finally, we consider a simple alternative model, in which the structural change terms are constructed slightly differently. We call the model that uses $MGROWTH$ and $DISEQ$, Model 1. To be implemented empirically, this model requires that $\phi = w_a L/Y$ is approximately constant across countries. We can relax this assumption, at the expense of assuming Cobb-Douglas technology in agriculture. If labor is paid its marginal product, then we have:

$$w_a = \mu \frac{Y_a}{L_a} = \mu \frac{sY}{aL}$$

where μ is the exponent on labor in the agricultural production function. Hence we have the following relationship:

$$\phi = \frac{w_a L}{Y} = \mu \frac{s}{a}$$

This suggests using the alternative set of explanatory variables in the growth regression:

$$\begin{aligned} MGROWTH2 &= (1-a)\frac{s}{a}\frac{\dot{m}}{m} \\ DISEQ2 &= \frac{p}{1-p}(1-a)\frac{s}{a}\frac{\dot{m}}{m} \end{aligned}$$

while ϕ is replaced by μ in the corresponding slope coefficients. The assumption that ϕ is the same across countries is replaced by an assumption that all countries have the same Cobb-Douglas technologies in agriculture (although TFP levels may differ). When we use $MGROWTH2$ and $DISEQ2$, we call this Model 2.

A major limitation of our empirical framework is worth noting: we have assumed that labor is homogeneous. Relaxing this assumption in a tractable way is not straightforward. In general, if workers differ in their skill levels, we would have to keep track of the composition of the workforce in each sector, and how it evolves as workers migrate. The need for this can be avoided, and our current approach justified, under the following assumptions. Workers of different skills are perfect substitutes at fixed ratios; workers face the same ratio of urban to rural wages, regardless of their skill level; and the skill composition is the same in both sectors - noting the necessary condition, that the skill distribution of migrants is always representative of that in each sector.

In this case, our earlier labor force variables L_a, L_m and L can be interpreted as measured in efficiency units of labor, and the non-agricultural employment share $m = L_m/L$ is the same whether expressed in terms of a simple head-count or in terms of efficiency units of labor. Our specification (12) continues to represent the output gains of reallocation. Intuitively, the skilled workers among the migrants can be treated 'as if' they are a larger number of unskilled workers, each with one unit of labor in efficiency-unit terms. Translating the reallocated workers into these units of unskilled labor, each unit of labor reallocated will raise output by the intersectoral difference in marginal products that holds for the unskilled, and hence the effect of reallocation is the one derived above.[12]

4 Dualism and structural change: stylized facts

This section describes the patterns of structural change observed in six regions of the world since 1960. Structural change has been substantial over this period, and the data are potentially consistent with significant wage differentials across sectors.

Table 1 shows figures for agriculture's share of employment (a) and share of nominal value added (s) for six regions, in 1960, 1980 and 1996. The figures are medians for each region, based on FAO and World Bank data.[13] Most regions of the developing world have seen a substantial change in sectoral structure over both 1960-80 and 1980-96. This can be measured in terms of an absolute change, or relative to the starting position. Based on the absolute change in the employment share, the shift out of agriculture appears to have been least

[12] The necessary assumptions are clearly highly restrictive. Temple (2005) describes another possible approach to resolving the problem, but in general, solutions are likely to require some combination of unattractively restrictive assumptions and data that are not readily available. See Graham and Temple (2006) for some additional discussion.

[13] The data on employment shares are from the Statistical Database of the Food and Agricultural Organization of the United Nations (FAO 2003). The data on value added shares are from the World Bank's (2002) World Development Indicators CD-Rom where available. Where necessary, the WDI data have been supplemented with figures for 1960 taken from the 1990 Production Yearbook of the FAO and the 1987 World Development Report of the World Bank.

pronounced in South Asia in 1960-80 and in sub-Saharan Africa in 1980-96. But when looking at the proportionate growth in non-agricultural employment, for 1960-80, this has been greatest in sub-Saharan Africa, rising from 12% of employment in 1960 to 24% in 1980. For 1980-96, it has been greatest in South Asia, rising from 30% to 40%.

Table 1 also reports a median figure for the ratio of the average product of labor in the two sectors, defined as:

$$RLP = \frac{qY_m/L_m}{Y_a/L_a} = \left(\frac{1-s}{s}\right)\left(\frac{a}{1-a}\right)$$

The table shows that average labor productivity is substantially higher outside agriculture, a well-known finding that is discussed in Kuznets (1971), Chenery and Syrquin (1975) and Gollin, Parente and Rogerson (2004), among others. This should not be used to conclude that agriculture, or factor allocation, is somehow inefficient. Differences in average products will usually be a feature of an efficient allocation, since output is maximized by equating marginal products rather than average products (see Temple 2005 for more discussion). Hence, data on average product differentials cannot establish the existence of dualism, without additional evidence or assumptions. We can be more confident of the following statement, however. If technologies are Cobb-Douglas with parameters that are roughly similar across the world, the rank ordering of marginal product differentials across regions will correspond to the rank ordering of the RLP figures. In other words, if there are significant marginal product differentials, it seems likely that they are greatest in sub-Saharan Africa. This will be taken into account in some of the empirical work that follows.

There are other reasons to be suspicious of the figures on relative productivity (and the data used to construct our structural change terms). It is likely that urban labor is more skilled on average, and that in poorer countries a substantial fraction of agricultural output is unmeasured in the national accounts, as discussed in Parente, Rogerson and Wright (2000). Schmitt (1989) points out the dangers of interpreting measures like RLP given that some agricultural labor is allocated to non-farm activities. For all these reasons, it seems likely that RLP overstates the relative productivity of workers in non-agriculture.

Another interesting aspect of Table 1 is that, for all regions but South Asia, RLP declines between 1960 and 1996. Based on earlier patterns, Chenery and Syrquin (1975, p. 53) argued that relative productivity in industry and services increases in the early stages of development, before ultimately declining. In contrast, the Table 1 figures suggest that the relative productivity of agriculture has improved even at low levels of development. This is consistent with a marginal product differential across sectors that is gradually being eliminated over time, although other explanations are also possible.

We now consider the pace of structural change in more depth, using the propensity to migrate as defined in equation (1). Table 2 shows the five countries with the most rapid structural change on this measure, and the five slowest. The general pattern is unsurprising: the countries with rapid changes include three well-known for fast growth (Japan, Korea and Singapore) while four of the countries with slow structural change are located in sub-Saharan Africa. This calls into question our earlier assumption that the speed of adjustment is similar across countries. Sub-Saharan African countries appear to be characterized by large marginal product differentials and slow reallocation. Our empirical work will sometimes use a specification in which the structural change terms are interacted with a dummy for sub-Saharan Africa. This allows the wage differential to be greater in Africa and/or the rate of adjustment to be slower.

5 Structural change and growth regressions

This section and the next will examine whether labor reallocation makes an important contribution to aggregate TFP growth. In this section, we add structural change terms to otherwise standard cross-country growth regressions, based on the specification of MRW. In the next section, we will estimate regressions in which measures of TFP growth, as computed by various authors, are used as the dependent variable. In both cases, we present results for a variety of specifications, and examine robustness in many dimensions. Our robustness checks include quantile regression and robust estimation, and restriction of the sample to developing countries. Given the possible concern that the extent of structural change will be endogenous in the technical sense (that is, correlated with the disturbance term in the regression) section 7 will present estimates from instrumental variable procedures.

5.1 Specification and data

We first discuss our treatment of human capital in the growth regressions. One of the main criticisms of the original MRW regressions has been their empirical treatment of human capital (see for example Gemmell 1996, Klenow and Rodriguez-Clare 1997 and Pritchett 2001). The human capital measure in MRW is based on the percentage of the working-age population that is in secondary school, obtained by multiplying the secondary enrollment rate by the fraction of the working-age population that is of school age. Consistent with MRW's theoretical derivation, this can be seen as a flow measure of the rate of investment in schooling.

Since direct measures of the stock of human capital are now available (Barro and Lee 2001) we integrate the level of human capital, rather than the rate of investment in schooling, into MRW's model. Our approach is based on equation (12) in MRW (p. 418), using data on the average level of human capital as a proxy for the steady-state stock. Note that this will alter the mapping between the slope coefficients and the underlying technology parameters. The human capital measure we use is average years of schooling in the population aged 15 and older, from Barro and Lee.

Our sample of countries is based on MRW's, which excludes oil producers and those with small populations. In their work, MRW looked at the time period 1960-85. We can now look at a longer time period, 1960-96, using the latest release of the Penn World Table, version 6.1 (Heston, Summers and Aten 2002). We have chosen 1996 as an endpoint because this maximizes the availability of data for the MRW set of countries. As well as considering 1960-96, we also work with two subperiods, 1960-80 and 1980-96, roughly corresponding to the periods before and after the onset of the debt crisis. Missing values in the Barro and Lee (2001) data set, or sometimes in PWT 6.1, force us to exclude a number of countries from the original MRW sample, so that we are left with a main sample of 76 developed and developing countries.

In the MRW regressions, the dependent variable will always be the log difference of output per worker over the relevant time period. One issue here is that the labor force may be mismeasured in PWT 6.1 for some countries. Dowrick (2005) notes some potential inconsistencies in the implied labor force participation rates for a number of African countries after 1980. Since 10 of these countries appear in our main 76-country sample, we follow Dowrick's recommendation and construct our own series for output per worker. We do this by combining the PWT 6.1 data on output with an alternative data set on the labor force, from the World Bank's World Development Indicators database.[14]

As a preliminary look at the patterns in the data, Table 3 reports correlations in the 76-country sample between the MRW variables and five new variables used in this paper. These are $MGROWTH$ and $MGROWTH2$, as calculated for 1960-96; $DISEQ$ and $DISEQ2$, the migration disequilibrium terms for the

[14]These data were downloaded in April 2006 and are taken from World Bank (2005). Our own comparisons between the PWT 6.1 and the World Bank labour force data confirmed the findings in Dowrick (2005). To see how the alternative labour force data modifies our results, the results in this paper can be compared with those in the working paper version, Temple and Woessmann (2005), which used the original PWT 6.1 data. The differences in results are generally minor.

same period; and a_{60}, the agricultural employment share in 1960.[15] As one might expect, $MGROWTH$ and $DISEQ$ are highly correlated, as are $MGROWTH2$ and $DISEQ2$.

The first column of Table 3 shows that the correlations of growth (DY) with the nonlinear structural change terms ($DISEQ$ and $DISEQ2$) are noticeably higher than the correlations with the linear structural change terms ($MGROWTH$ and $MGROWTH2$). This is preliminary support for our new specification, which implies that the cross-section relationship between growth and the extent of structural change should be convex, rather than linear.

There is some further evidence for this convex relationship in the data, at least when we condition on the other explanatory variables. Figure 1 shows an added-variable (partial scatter) plot of growth over 1960-96, conditional on four explanatory variables (investment, human capital, population growth and initial income) and regional dummies, against $MGROWTH$ conditional on the same variables. A quadratic regression line added to the plot suggests there is indeed some convexity in the growth-structural change relationship. The growth impact of a given extent of structural change appears to be greatest in those countries experiencing more rapid structural change. The graph shown here is for the main sample of 76 countries, but the convexity is similarly evident in a plot restricted to 56 developing countries (not shown). Our subsequent regressions will investigate the relationship in more detail. Initially we will focus on the precision of the estimates, before discussing their magnitude in section 8.

5.2 Initial evidence

We begin by estimating the standard MRW specification for 1960-85. The dependent variable is the log difference of output per worker between 1960 and 1985, and the explanatory variables are the log of the average investment share, the log of a measure of schooling investment, the log of the average population growth rate plus 0.05, and the log of initial income. For all our OLS regressions, the estimated standard errors are corrected for heteroscedasticity using the method of White (1980).

The first results are presented in Table 4. For comparison with the original MRW results, regression (1) shows their model re-estimated using the revised PWT 6.1 data for output and the WDI data for the labor force. This excludes 6 of MRW's original sample of 98 countries. The results closely resemble their Table V findings. In regression (2), we estimate their model for our main sample

[15] In constructing the structural change terms, we use the initial employment share a_{60} as the number corresponding to a in their definitions.

of 76 countries, obtaining very similar results.

We now compare their results to our alternative specification for human capital, supplement the specification with our structural change terms, and extend the period to 1996. First of all, we consider the alternative MRW specification with human capital levels, shown in regression (3). This replaces their human capital measure ($SCHOOL$) with the logarithm of average years of schooling in the working-age population averaged over the time period. The coefficients are not directly comparable to the MRW specification, but continue to provide support for the effects implied by the augmented Solow model. The estimates imply an output-capital elasticity of 0.57, an output-human capital elasticity of 0.41, and a convergence rate of 1.2% a year. These values are of the same order of magnitude as those obtained by MRW (their Table VI) although our estimates of the output elasticities are higher than in MRW.

Regression (4) supplements the model with the structural change terms, $MGROWTH$ and $DISEQ$ calculated for 1960-85. They are not individually significant, but the nonlinear term $DISEQ$ is approaching significance at the 10% level and there is strong evidence of joint significance, as revealed by the corresponding Wald test (F-statistic 28.20; p-value 0.00). Hence the MRW specification is firmly rejected in favour of the more general specification that we adopt. Moreover, the inclusion of the structural change terms raises the explanatory power of the growth regression to an unusual extent. The R^2 rises from 0.43 to 0.59.

We now move to a longer time period, 1960-96, shown in regression (5). The results are similar. The parameter values implicit in the coefficient estimates, after adjusting for the altered length of the time period, are largely unchanged. $DISEQ$ is now significant at the 10% level, when both structural change terms are included. We also consider whether the structural change terms are significant when entered on their own. We refer to these as restricted models, and report the associated results as regressions (6) and (7). In either case, the single structural change term is significant at the 5% level. The model based on the nonlinear term $DISEQ$ has slightly greater explanatory power.

Finally, regressions (8) and (9) are based on two subperiods, 1960-80 and 1980-96. In the first subperiod, the two structural change terms are positively signed and jointly significant at the 1% level. The results are noticeably weaker for 1980-96, although the two terms are jointly significant at the 15% level. It is possible the weaker results reflect a decline in the extent of dualism over the course of the 1980s and early 1990s.

5.3 Further evidence

We now consider further evidence, presented in Table 5. The time period is 1960-96 throughout. All the regressions from this point onwards include four regional dummies, corresponding to sub-Saharan Africa; non-OECD East Asia and the Pacific; Latin America and the Caribbean; and the high-income OECD countries, using the World Bank (2002) classifications. The coefficients on the regional dummies are not reported.

Regressions (10)-(12) in Table 5 show that our earlier findings are robust to the inclusion of regional dummies. The structural change terms are jointly significant (regression 10) or individually significant in the restricted models (regressions 11 and 12). $MGROWTH$ is negatively signed in regression (10), but the estimates are imprecise given the use of both structural change terms. The disequilibrium term $DISEQ$ dominates in this specification, and is significant even when both structural change terms are included.

In regression (13), we add the initial share of employment in agriculture ($a60$) as an explanatory variable. This allows us to check that the structural change terms are not simply a proxy for initial specialization in agriculture, which could affect growth for a wide variety of reasons.[16] As shown in regression (13), allowing for this effect does not change the results. The new variable is not significant at conventional levels and the structural change terms remain jointly significant (F-statistic 12.1; p-value 0.00).

In section 3, we also derived an alternative specification for the structural change terms, which we called Model 2. In regressions (14)-(16) we show that our findings are robust to this alternative specification. This is the case even though we now have fewer observations, due to lack of the necessary data on agriculture's share of value added. $MGROWTH2$ and $DISEQ2$ are jointly significant (F-statistic 7.94, p-value 0.00) and are significant when entered separately.

What can we conclude thus far? There is clear evidence for structural change effects associated with marginal product differentials. The two structural change terms almost always have the predicted signs, are jointly significant, and greatly increase the explanatory power of otherwise standard growth regressions. The $DISEQ$ variable performs especially well, and this suggests that empirical models should allow the relationship between growth and structural change to be nonlinear. There is also some tentative evidence that the extent of dualism,

[16] An alternative approach would be to use the initial value added share of agriculture, or an average of that share over time, or even the average share interacted with regional dummies. Our experiments with these various alternatives indicated that the structural change terms remain jointly significant. We prefer to use the employment share because it is available for a larger number of countries.

as reflected in marginal product differentials, has declined over time. We will explore this further in section 8.

5.4 Robust and quantile regressions

We now perform several further robustness checks, mainly to ensure that our results are not driven by outlying observations. The message of these tests is that our results are unusually robust. The results are contained in Table A1 in the appendix, and the discussion that follows could be skipped by readers more interested in our overall conclusions than in the details of robustness tests.

First, we estimate the regressions with the median regression estimator (MR), also known as the LAD estimator. This estimator minimizes the sum of the absolute residuals, and is therefore less sensitive to outliers than an estimator like OLS that minimizes the sum of squares.[17] The results can be found in Appendix Table A1, as regressions (A1)-(A6). Our findings are generally robust to the use of median regression. The two structural change terms are not jointly significant in Model 2, but are each significant at the 1% level when entered individually.

Second, we also implement an alternative robust regression technique that drops or downweights outliers. The method we use starts by eliminating gross outliers for which Cook's distance measure is greater than one, and then iteratively downweights observations with large absolute residuals.[18] Our results are essentially unchanged on downweighting these observations; see regressions (A7) and (A8).

We have also estimated regressions that exclude Singapore (see regression A9). This country combines fast growth with by far the highest value of the propensity for migration, as listed in Table 2. This is likely to reflect a very small agricultural sector, given that Singapore is a city-state. The OLS, MR and RR results are all robust to the exclusion of Singapore. One difference is that the nonlinear term is no longer individually significant in the unrestricted model containing both terms. Nevertheless, the structural change terms retain joint significance at the 1% level.

Overall, we conclude that our results are not driven by the presence of outliers. We now consider whether the effects of structural change vary across

[17] In particular, the MR estimator may be preferable to least squares when the distribution of the regression errors has thick tails. If the errors are i.i.d. with a Laplace distribution, and distributed independently of the explanatory variables, then the MR estimate of a linear model is also the maximum likelihood estimate.

[18] This estimator corresponds to the *rreg* robust estimation command in Stata.

regression quantiles.[19] Figure 2 plots the 9 quantile regression estimates for each 0.10 percentile interval for the two restricted models. For either structural change term, the estimated effect is broadly similar across the full range of quantiles of the conditional growth distribution, and almost always lies within the confidence interval of the OLS estimate (the exception here is the 80th percentile estimate for $DISEQ$).[20] The interpretation of this result depends on the sources of the disturbances. The tendency to a downwards slope suggests that the effects of structural change are slightly smaller in economies at the upper end of the conditional distribution.

5.5 Structural change in developing countries

Our empirical work thus far has used a large sample of countries, developed and developing, with very different sectoral structures and patterns of structural change. We now examine whether the structural change effects can be identified even in a sample restricted to developing countries. To achieve this, we exclude the 20 countries in our main sample that are classified as high-income OECD countries in World Bank (2002). This set of high-income countries is broadly the same as the group of OECD members in the late 1960s, and hence excluding them should leave us with a sample that corresponds reasonably well to those countries considered less developed in the 1960s. The final three columns in Table 5 are based on this restricted sample, and regressions (17)-(19) show that our previous findings apply even when developed countries are excluded. We also consider robustness issues for this sample (56 countries for Model 1 and 48 countries for Model 2). The results are shown in Appendix Table A2, are qualitatively the same as before, and the next two paragraphs can be skipped by readers more interested in the overall findings.

Earlier in the paper, we noted the possibility of slower adjustment in sub-Saharan Africa. We have estimated regressions which include interaction terms, in which $MGROWTH$ and $DISEQ$ are interacted with the Africa dummy. These interaction terms are statistically insignificant, even jointly. The nonlinear term $DISEQ$ remains significant (regression A13 in Appendix Table A2). These results are tentative evidence that a simple model may capture the growth effects of structural change adequately even for sub-Saharan Africa.

We have also carried out some robustness tests for the developing country sample, based on MR estimation. The two structural change terms are jointly

[19] See Koenker and Hallock (2001) for an introduction to quantile regression.

[20] Note that in a sample of this size, there is likely to be considerable uncertainty associated with estimating the relationship that holds at the extremes of the conditional distribution.

significant only at the 15% level, but are each significant at 10% when entered individually. In further results, not reported, we have confirmed that our findings are not sensitive to outliers and the use of robust estimation. This includes the results for the Model 2 specification based on $MGROWTH2$ and $DISEQ2$.

6 Structural change and TFP growth

Thus far, our examination of structural change and growth has been based on cross-section growth regressions, with all their attendant econometric problems. Some of the problems, such as the possible endogeneity of the investment variable, may cause the impact of structural change to be understated. Nevertheless, it is also possible that the regressions overstate the extent of structural change effects.

We now consider regressions in which the dependent variable is a measure of TFP growth constructed by previous researchers, using growth accounting. We have used three measures: primarily estimates of TFP growth rates over 1960-85 due to Klenow and Rodriguez-Clare (1997), but also estimates for 1965-95 due to Bernanke and Gürkaynak (2001) and for 1960-2000 due to Bosworth and Collins (2003).[21] In principle, we should measure TFP growth for each country using country-specific factor shares, but these are hard to obtain for developing countries, especially given problems raised by self-employment and unincorporated enterprises (Gollin 2002). For this reason, we use TFP growth rates that assume common factor shares across countries.[22] Each TFP series includes an adjustment for the growth of human capital.

The sample is again based on the non-oil set of countries used by MRW, but limited by data availability. Note that the coefficient estimates will not be directly comparable with earlier results. This is because our earlier growth regressions are based (as in MRW) on the log difference of GDP per worker over the respective periods, whereas the TFP growth regressions use the compound growth rate of TFP measured in annual percentage points. All our TFP regressions include the same set of regional dummies used previously. Since TFP growth is likely to reflect, at least in part, a process of technological catch-up, we have included the log of initial TFP as an additional control variable whenever it can be constructed from the available data.

We begin with the Klenow and Rodriguez-Clare measure (KR). The Table 6

[21] We are grateful to these authors for making their calculations of TFP growth available.

[22] Bernanke and Gürkaynak (2001) calculate alternative TFP growth estimates using data on country-specific factor shares, but for a much smaller sample of countries. Our results are robust to using these alternative data (not shown).

results for this measure, regressions (20)-(25), tell a story very similar to that of the previous regressions. In the unrestricted model for the full sample, reported as regression (20), the two structural change terms $MGROWTH$ and $DISEQ$ are jointly significant (F-statistic 19.61, p-value 0.00). When entered separately, as in the restricted models shown in regressions (21) and (22), each term is positive and significant at the 1% level. Results for the Model 2 specification, based on $MGROWTH2$ and $DISEQ2$, are very similar, as can be seen from regressions (23)-(25). Once again the two structural change terms are jointly significant. Also note that these models account for around half the international variation in TFP growth. For comparison, when the structural change terms are dropped the R^2 falls to 0.33.

When the initial agricultural employment share is included as an additional control variable, it is rarely significant in these TFP growth regressions, and it never changes the results on the structural change terms. Our findings continue to be robust to the use of MR and robust regression, and to the exclusion of Singapore. Results are similar when the sample of countries is restricted to the developing country sample (as in regressions A21 and A22 in Appendix Table A3). As before, the results are robust to including interactions of the structural change terms with a dummy for sub-Saharan Africa.

We now consider alternative measures of TFP growth. Bernanke and Gürkaynak (2001) provide a range of TFP growth measures for 1965-95. We focus on the series which assumes a labor share of 0.65 and an annual return to additional years of schooling of 7%. We construct structural change terms for 1965-95, but the FAO (2003) data on employment shares are available only for 1960 and 1970. We approximate the 1965 value by a mean of the two. The levels of statistical significance are slightly lower when using this TFP growth measure (regressions 26-28) but the general pattern of results is not greatly different from previous findings. When using the Bosworth and Collins measure for 1960-2000, the results are weaker, but Model 2 has some explanatory power (regression 29). Note that we cannot control for the initial level of TFP in these regressions because Bosworth and Collins construct their TFP series using national prices.

Overall our results are consistent with the hypothesis that the reallocation of labor, in the presence of marginal product differentials, makes a sizeable contribution to the aggregate Solow residual. Our relatively simple models explain a significant fraction of the observed residual, and therefore help to chip away at this 'measure of our ignorance'.

7 Instrumental variable estimates

Structural change is clearly an endogenous process, driven by a variety of economic forces. Whenever the relationship between TFP growth and structural change is estimated from the data, a major concern is that the extent of structural change may be endogenous also in the statistical sense, namely correlated with the regression disturbances. Informally, one might expect the coefficients on structural change terms to be biased away from zero. The magnitude of this effect is an open question, but here we attempt to address the problem using instrumental variable methods, including 2SLS, GMM and Fuller's (1977) modification of the LIML estimator. The motivation for using Fuller's estimator is the weakness of our instruments, something that we discuss further below.

In the present context, the main candidates for instruments will be variables that affect either the potential supply of migrants, or the incentives to migrate, or both. Our primary instrument is $POPAGE$, the share of the population aged between 0 and 14 in 1960. The young are likely to have particularly strong incentives to migrate and, over the course of our time period, those aged between 0 and 14 in 1960 will have reached the 15-30 age group among which migration tends to be concentrated (Mazumdar 1987, p. 1119). We also use the log of relative labor productivity, $LRLP$, at the beginning of the period, since the incentives to migrate may be correlated with the observed extent of dualism. Relative labor productivity is defined as in section 4.

Although both these instruments have some appeal, it would be easy to criticise the associated exclusion restrictions. We also experiment with an alternative strategy, which is to estimate growth regressions for 1980-96 and include lagged structural change terms (1960-80) among the instruments.

When using any of these approaches, estimating the full model with both structural change terms is ambitious. It requires us to find an instrument set such that the fitted values of the endogenous explanatory variables are not highly correlated, and this is difficult to achieve in practice. For this reason, we focus on restricted models (with just one structural change term) throughout.

All our IV models contain regional dummies and the MRW regressors, but we only report the coefficient and standard error on the structural change term, for ease of comparison across the different estimation methods, including OLS for comparison.[23] Our various experiments with an IV strategy were more successful for Model 2, and so we concentrate on that specification. We first look at 1960-96, and report the results in regressions (30) and (31) of Table 7. The coefficients

[23] To implement the different estimators, we use the `ivreg2` software for Stata. See Baum et al. (2003).

on either structural change term, $MGROWTH2$ or $DISEQ2$, are significantly different from zero at the 5% level.

Based on the 2SLS results, we implement a Wu-Hausman test. This does not reject the exogeneity of the structural change terms at conventional levels, but comes close to doing so (p-value of 0.26 for $MGROWTH2$ and 0.19 for $DISEQ2$). It is important to note, however, that this near-rejection does not arise because the 2SLS estimates of the structural change parameters are closer to zero than before (the expected pattern). Instead, the 2SLS coefficients are larger than the OLS estimates. This pattern tends to suggest either that the exclusion restriction is invalid, or that measurement error in the structural change terms has led to an attenuation bias in the OLS coefficient estimates. Although the overidentifying restrictions are not rejected by a Sargan test (for 2SLS) or Hansen's J-test (for GMM) there must be a significant question mark over the exogeneity of the instruments, given the pattern of the coefficient estimates.

Finally, we consider estimates of Model 2 for the 1980-96 period (where Model 1 again works less well). Here we instrument using $POPAGE$ and the lagged value of the structural change term, calculated over 1960-80. In these estimates, $MGROWTH2$ is not significant at conventional levels using IV methods, but $DISEQ2$ performs more strongly. Again we do not reject the validity of the overidentifying restrictions, and the Wu-Hausman tests find no evidence of endogeneity of the structural change terms.

We now discuss the strength of the instruments. As is now well known, when instruments are only weakly correlated with the endogenous explanatory variables, the 2SLS and GMM estimators may be badly biased in small samples. Moreover, the conventional asymptotic approximations used for hypothesis tests and confidence intervals are likely to be unreliable. Studies such as Stock, Wright and Yogo (2002) have suggested, as a rule of thumb, that values for the first-stage F-statistic below 10 indicate a weak instrument problem. Three of our first-stage F-statistics are below this threshold, and for this reason we have also reported estimates based on Fuller's (1977) modification of the LIML estimator. Fuller's estimator may be more robust than 2SLS in the presence of weak instruments, and is designed to ensure the estimator has finite moments (unlike LIML). It performs relatively well in the simulations carried out in Hahn, Hausman and Kuersteiner (2004) and appears to have lower small-sample variability than LIML. We set the user-specified constant (denoted by alpha in Fuller 1977) to a value of one, at which point the estimator is nearly unbiased (Fuller 1977, p. 951).[24] It can be seen from Table 7 that the use of the Fuller (1) estimator

[24]The point estimates tend to be slightly smaller in magnitude if we set alpha to four (a

gives results broadly comparable to those obtained with 2SLS and GMM.

An alternative response to weak instrument biases is the use of robust methods for inference, such as those developed by Moreira (2003) and recommended by Stock, Wright and Yogo (2002). Given the potential weakness of our instruments, we have used Moreira's method to construct robust 95% confidence regions for the structural change coefficients, and these are also reported in Table 7. We base these confidence regions on conditional likelihood ratio tests, which appear to have good power properties (Moreira 2003). As Table 7 shows, these confidence regions tend to be wide, and always include zero; for regression (31) the confidence region is so wide that it is not well-defined.

In summary, the message from the IV results is mixed. On the positive side, we can sometimes obtain reasonably precise estimates of the coefficients on the structural change terms. In common with several other applications of IV methods to growth questions using cross-country data, we find that the IV estimates are not only significant, but further away from zero than the OLS estimates. At first glance, the differences between the OLS and IV coefficients suggest that the expected simultaneity bias is either not present, or has been offset by other factors such as measurement error.

At the same time, there is a need for caution. The tendency for higher coefficients under IV suggests that the exclusion restrictions may be questionable, and this is a particular concern given the weakness of our instruments, as the two problems will tend to reinforce each other. Moreover, the coefficients on the structural change terms are imprecisely estimated under a range of alternative specifications. In particular we find much weaker results (not reported) when using TFP growth as the dependent variable. Then, either the models are only weakly identified, or the standard errors on the structural change terms are too high to draw useful conclusions about the parameters. In samples of this size, all the coefficient estimates and specification tests may be sensitive to small numbers of observations, but there is no generally agreed-upon method to ensure robustness in the IV context. For all these reasons, we are inclined to place more weight on the OLS findings earlier in the paper. But we must also acknowledge the possible endogeneity of structural change as a key drawback of those results, a weakness shared with previous studies.

value which may improve the performance of the estimator in mean-square error terms) but the reductions are not large enough to modify our overall conclusions.

8 The implied parameter values

So far, we have shown that structural change terms have some explanatory power when included in either standard growth regressions or TFP growth regressions. In this section, we focus on the magnitude of the associated parameter estimates, rather than simply their precision. We calculate the parameter values implied by the OLS results, based on transformations of the regression coefficients, and also obtain an alternative set of parameter estimates more directly, by using nonlinear least squares (NLS) estimation.

We are able to show that our regression estimates imply marginal product differentials of a similar magnitude to the rural-urban wage differentials sometimes observed in microeconomic data. Moreover, since the nonlinear term allows the estimated differential to vary across countries, we calculate and report the extent of this variation. This is not only of independent interest, but also acts as a check that our regression specification and parameter estimates do not have implausible implications.

First of all, we briefly discuss the microeconomic evidence on rural-urban wage differentials. This evidence is patchy, with reliable data available for only a small number of countries. The data in World Bank (1995, p. 76) suggest that the urban wage can easily be 30-100% higher than the rural wage for workers of similar skill levels. As we noted in the introduction, however, wages may depart from marginal products, for example because workers receive their average product in the agricultural sector (Lewis 1954). In this case, marginal product differentials could be much larger than observed wage gaps. Our estimation of a production relationship allows the extent of differentials to be inferred for a large number of countries, at the expense of some strong assumptions.[25]

Using our theoretical model, and a small number of parameter assumptions, the coefficients in our regressions can be used to calculate the values of the parameters in the model. First of all, we focus on obtaining an estimate of k, the fixed wage differential that arises when $DISEQ$ is excluded. We then look at the nonlinear model, which sets $k = 1$ but allows the implicit wage differential to vary when $p \neq 0$.

There are a few technical issues here that could be skipped by readers interested primarily in the final results and their economic interpretation. First of all,

[25] It is the marginal product differential, not the wage gap, that will drive our empirical results. The marginal product differential determines the effects of structural change on productivity growth, and thereby influences the partial correlations between structural change and growth observed in the data. In the remainder of this section, we will use the term 'wage differential' as a convenient shorthand, but our estimates are best seen as relating to the magnitude of the marginal product differential.

our model is set up in such a way that structural change influences TFP growth. In our MRW-style growth regressions, we have to rescale the coefficients so that they correspond to effects on annual TFP growth rather than overall growth in labor-augmenting efficiency. This is easily done, and we denote the rescaled coefficient on $MGROWTH$ as π; this rescaled coefficient correspond to $(k-1)\phi$ in our model. In order to calculate the implied k, we need an assumption about $\phi = w_a L/Y$. This parameter will be close to the aggregate labor share if the agricultural sector accounts for the majority of employment and/or the inter-sectoral wage gap is not large. We adopt a value of 2/3 for ϕ, but the order of magnitude of the implied differential does not hinge on the assumption about ϕ, and our results would not be greatly changed by considering $\phi = 1/2$.[26]

Table 8 presents the parameter values implicit in our growth regressions and TFP regressions. The first case is the restricted model without $DISEQ$. The calculation can be illustrated with an example. In the case of regression (11) in Table 5, the model yields a coefficient estimate on $MGROWTH$ of 0.62. Dividing by the number of years, given that the dependent variable is the log difference of output between 1960 and 1996, and rescaling by one third (to get from labor-augmenting efficiency to TFP growth) and multiplying by 100 (due to the scaling of the variable) yields $\pi = 0.58$. This implies a value of $k = 1.87$. That is, the marginal product of labor in non-agriculture is close to double that in agriculture. Across a wide range of models, samples and estimation methods, the implied marginal product ratio lies between 1.8 and 3.8.

A limitation is that, within a given regression, the wage differential is assumed to be constant across countries. It is therefore interesting to explore the model with only $DISEQ$ and assuming $k = 1$. The theoretical model implies that the coefficient on $DISEQ$ is equal to ϕ/ψ. The coefficient on $DISEQ$ from regression (12) in Table 5 was 1.76. After rescaling, this implies a value for the speed of adjustment parameter ψ of 0.041. We can interpret this as follows. In our main sample, the median propensity to migrate (p) is 0.0199. Using equation (4) this implies a current wage ratio w_m/w_a of 1.50 for a country with the median value of p. In this specification, however, the implicit wage ratio varies across countries. For the country at the 10th percentile of the p distribution, the implied wage ratio is 1.08, while it is 2.06 at the 90th percentile. Alternative specifications give similar results. These results are promising in that they indicate low marginal product differentials in some countries (associated with

[26] Such assumptions may not be unreasonable in the light of Gollin (2002). He argues that the aggregate labour share is not systematically related to the level of development, although it does vary across countries. See Durlauf's comments on Bosworth and Collins (2003) for more discussion.

a low propensity to migrate) while in others, the marginal product of labor is substantially higher in non-agriculture than in agriculture.

Another approach is to use nonlinear least squares to estimate the parameters directly. This allows us to replace our assumption that ϕ is constant across countries with an assumption that the aggregate labor share η is constant across countries. This is done by substituting ϕ out of the regression equation, using equation (19) in the Appendix. In the NLS regressions, we assume the aggregate labor share $\eta = 2/3$. The results are shown in Appendix Table A4, and yield parameter values that are in line with those reported above, although somewhat higher for the TFP growth regressions.

In summary, a wide variety of specifications and estimation methods combine to tell a plausible story. The nonlinear model, in particular, implies that marginal product differentials are of a similar order of magnitude to those found in microeconomic studies, but are barely present in a subset of economies, namely those where recent structural change has been limited. Another finding, made clear by comparing across the last four columns of Table 8, is that the implied magnitude of the differentials was noticeably lower in 1980-96 than in 1960-80. This is consistent with the view that the extent of dualism has declined over time.

9 Summary and conclusions

Current empirical growth models are often criticised for neglecting structural change. When there is a differential in the marginal product of labor across sectors, changes in the structure of employment will be an independent source of growth in aggregate productivity. This paper presents an empirical growth model that incorporates this effect and allows marginal product differentials to vary across countries in a more flexible way than previous work. Our regressions quantify the productivity effects of structural change for the period 1960-96.

We find sizeable differentials, comparable in magnitude to microeconomic evidence. There is some evidence of variation across countries, and we also find that the differentials have fallen over time. The results support the emphasis of Vollrath (2005) on factor market distortions and their consequences for aggregate productivity. Consistent with his static analysis, we find that structural change can account for a significant fraction of the observed variation in productivity growth. Our regressions for productivity growth that include regional dummies, the initial level of TFP, and structural change terms can explain around half the cross-section variation. When structural change terms are excluded, this proportion falls to a third.

The frameworks developed here could be extended in several ways. Above all, it would be interesting to consider other testable implications of dual economy models. Temple (2005) discusses some of the relevant issues. One obvious modification to the analysis above would be to assume that agricultural labor receives its average product rather than its marginal product. It would also be interesting to explore whether our empirical findings arise because of genuine marginal product differentials, or because of sector-specific production externalities of the form considered in Graham and Temple (2006). More broadly, it seems likely that empirical growth research could benefit from closer engagement with the dual economy tradition, and there are many opportunities for further work in this direction.

10 Appendix

10.1 Relative prices

First of all, we discuss how our analysis can incorporate the effects of time-varying relative prices. Our approach to price variation follows Jorgenson and Griliches (1967), also briefly discussed in Barro and Sala-i-Martin (2004, p. 443, footnote 6). The analysis is based on modelling changes in real output, using a GDP price index. This price index should be regarded as distinct from a cost-of-living index since, given the small open economy setting, the structure of consumption will typically differ from that of production.

The first step is to differentiate (5) with respect to time. This gives an expression for growth in real output:

$$\frac{\dot{Y}}{Y} = s(t)\frac{\dot{Y}_a}{Y_a} + (1 - s(t))\frac{\dot{Y}_m}{Y_m} + (1 - s(t))\frac{\dot{q}}{q} - \frac{\dot{\Omega}}{\Omega} \tag{14}$$

If we use a Divisia price index for the output deflator, the growth rate of this price index is equal to a share-weighted average of the growth rates of the component prices. In our case, given that the agricultural good is the numeraire, we have:

$$\frac{\dot{\Omega}}{\Omega} = (1 - s(t))\frac{\dot{q}}{q}$$

Hence (14) leads to the expression given in the main text,

$$\frac{\dot{Y}}{Y} = s(t)\frac{\dot{Y}_a}{Y_a} + (1 - s(t))\frac{\dot{Y}_m}{Y_m} \tag{15}$$

which says that real output growth is a share-weighted average of the growth in quantities (corresponding to a Divisia output index). The same logic carries over to the case with a marginal product differential, although the formal justification for using Divisia indices is weaker here, given the presence of a distortion.

This analysis implies that movements in prices do not enter our expressions for the aggregate Solow residual directly, but only indirectly, as components of the value added shares. Movements in prices may affect the Solow residual indirectly by changing the extent of migration, or through the value added shares; but conditional on these effects, there should be no *direct* effect of price variation on the Solow residual.

Our results have further implications: our empirical analysis is consistent with relative prices that differ across countries (due to tariffs or export subsidies, say). In effect, the role of relative price variation in determining domestic allocations is captured entirely in the output share data. A closely related point is that our empirical framework is developed entirely in terms of these nominal shares of value added, without any kind of PPP-type adjustment for price differences across countries. Again, this makes good economic sense. The domestic allocation of factors across sectors depends on the relative prices faced by domestic producers: *conditional on these prices*, the fact that producers in other countries may face a different set of prices does not affect the domestic production equilibrium, including the allocation of factors across sectors. In other words, as our derivations imply, the appropriate choice in this context is to use nominal shares of value added. By definition, these are computed using current domestic prices rather than prices that are constant across time or space. The use of PPP adjustments would be essential if we attempted to compare sectoral productivity *levels* across countries, but our analysis makes no use of such comparisons.

10.2 Derivation of the expression for the Solow residual

We now describe the derivation of equation (12) in the main text. First of all, given the agricultural production function, agricultural output growth is equal to:

$$\frac{\dot{Y}_a}{Y_a} = \frac{\dot{A}_a}{A_a}\frac{A_a F(.)}{Y_a} + \frac{A_a F_K K}{Y_a}\frac{\dot{K}_a}{K} + \frac{A_a F_L L}{Y_a}\frac{\dot{L}_a}{L}$$

Using (7), (8), the definition of s and the definitions of factor shares, we can write:

$$s\frac{\dot{Y}_a}{Y_a} = s\frac{\dot{A}_a}{A_a} + (1-\eta)\frac{\dot{K}_a}{K} + \phi\frac{\dot{L}_a}{L} \tag{16}$$

where we suppress time subscripts for notational simplicity. Output growth in non-agriculture is (in real terms):

$$\frac{\dot{Y}_m}{Y_m} = \frac{\dot{A}_m}{A_m}\frac{A_m G(.)}{Y_m} + \frac{qA_m G_K K}{qY_m}\frac{\dot{K}_m}{K} + \frac{qA_m G_L L}{qY_m}\frac{\dot{L}_m}{L}$$

Hence we can write:

$$\begin{aligned}(1-s)\frac{\dot{Y}_m}{Y_m} &= (1-s)\frac{\dot{A}_m}{A_m} + (1-\eta)\frac{\dot{K}_m}{K} + \frac{w_m L}{Y}\frac{\dot{L}_m}{L} \\ &= (1-s)\frac{\dot{A}_m}{A_m} + (1-\eta)\frac{\dot{K}_m}{K} + \phi\frac{\dot{L}_m}{L} + \frac{w_m L}{Y}\frac{\dot{L}_m}{L} - \phi\frac{\dot{L}_m}{L}\end{aligned}$$

Now we can use our expression for the modern sector wage (4) to obtain:

$$\begin{aligned}(1-s)\frac{\dot{Y}_m}{Y_m} &= (1-s)\frac{\dot{A}_m}{A_m} + (1-\eta)\frac{\dot{K}_m}{K} + \phi\frac{\dot{L}_m}{L} \\ &\quad + \frac{kw_a L}{Y}\frac{L_m}{L}\frac{\dot{L}_m}{L_m} + \frac{1}{\psi}\frac{p}{1-p}\frac{kw_a L}{Y}\frac{L_m}{L}\frac{\dot{L}_m}{L_m} - \phi\frac{L_m}{L}\frac{\dot{L}_m}{L_m}\end{aligned}$$

Using $\phi = w_a L/Y$ and $L_m/L = 1-a$ we can rewrite this as:

$$\begin{aligned}(1-s)\frac{\dot{Y}_m}{Y_m} &= (1-s)\frac{\dot{A}_m}{A_m} + (1-\eta)\frac{\dot{K}_m}{K} + \phi\frac{\dot{L}_m}{L} \\ &\quad + (k-1)\,\phi(1-a)\frac{\dot{L}_m}{L_m} + k\phi(1-a)\frac{1}{\psi}\frac{p}{1-p}\frac{\dot{L}_m}{L_m}\end{aligned} \tag{17}$$

We can combine (15), (16) and (17) and use $\dot{K}_a + \dot{K}_m = \dot{K}$ and $\dot{L}_a + \dot{L}_m = \dot{L}$ to obtain an equation for aggregate growth:

$$\begin{aligned}\frac{\dot{Y}}{Y} &= (1-\eta)\frac{\dot{K}}{K} + \phi\frac{\dot{L}}{L} + s\frac{\dot{A}_a}{A_a} + (1-s)\frac{\dot{A}_m}{A_m} + \\ &\quad + (k-1)\,\phi(1-a)\frac{\dot{L}_m}{L_m} + k\phi\frac{1}{\psi}\frac{p}{1-p}(1-a)\frac{\dot{L}_m}{L_m}\end{aligned} \tag{18}$$

We can simplify this further, as follows. The aggregate labor share is:

$$\eta = \frac{w_a L_a + w_m L_m}{Y}$$

Equation (4), together with $L_a = aL$ and $L_m = (1-a)L$ implies:

$$\eta = \frac{w_a aL + \left[kw_a + \frac{1}{\psi}\frac{p}{1-p}kw_a\right](1-a)L}{Y}$$

Using $\phi = w_a L/Y$ we have

$$\begin{aligned}\eta &= \phi\left[a + (1-a)k + \frac{1}{\psi}\frac{p}{1-p}(1-a)k\right] \\ &= \phi\left[1 + (1-a)(k-1) + \frac{1}{\psi}\frac{p}{1-p}(1-a)k\right]\end{aligned}$$

Hence

$$\eta - \phi = \phi \left[(1-a)(k-1) + \frac{1}{\psi} \frac{p}{1-p} (1-a) k \right] \tag{19}$$

Also note that we can write:

$$\phi \frac{\dot{L}}{L} = \eta \frac{\dot{L}}{L} - (\eta - \phi) \frac{\dot{L}}{L}$$

Using this relationship, and (19) together with (18) implies that aggregate growth equals

$$\begin{aligned} \frac{\dot{Y}}{Y} &= (1-\eta) \frac{\dot{K}}{K} + \eta \frac{\dot{L}}{L} + s \frac{\dot{A}_a}{A_a} + (1-s) \frac{\dot{A}_m}{A_m} + \\ & + (k-1)\, \phi (1-a) \left[\frac{\dot{L}_m}{L_m} - \frac{\dot{L}}{L} \right] + k\phi \frac{1}{\psi} \frac{p}{1-p} (1-a) \left[\frac{\dot{L}_m}{L_m} - \frac{\dot{L}}{L} \right] \end{aligned} \tag{20}$$

If we define $m = L_m/L = 1-a$ then the growth equation can be rewritten as:

$$\begin{aligned} \frac{\dot{Y}}{Y} &= (1-\eta) \frac{\dot{K}}{K} + \eta \frac{\dot{L}}{L} + s \frac{\dot{A}_a}{A_a} + (1-s) \frac{\dot{A}_m}{A_m} + \\ & + (k-1)\, \phi (1-a) \frac{\dot{m}}{m} + k\phi \frac{1}{\psi} \frac{p}{1-p} (1-a) \frac{\dot{m}}{m} \end{aligned} \tag{21}$$

which then implies equation (12) in the main text.

10.3 Approximations used in estimating the model

We now describe some of the assumptions involved in implementing the model empirically. As described in the text, we can test the model using regressions with aggregate TFP growth as the dependent variable. But some of our empirical work proceeds by adding structural change terms to growth regressions of the MRW form. The theoretical derivation in MRW, which leads to a linear regression, is developed for a one-sector model with a Cobb-Douglas production function:

$$Y = K^{\alpha} H^{\beta} (AL)^{1-\alpha-\beta}$$

where the notation is standard. MRW then show that the change in log output per worker between periods 0 and t can be approximated by

$$\log \frac{Y(t)}{L(t)} - \log \frac{Y(0)}{L(0)} = \theta \log A(0) + gt + \theta \gamma' X - \theta \log \frac{Y(0)}{L(0)}$$

where $A(0)$ is the initial level of labor-augmenting efficiency, g is the growth rate of efficiency A, θ is a parameter related to the convergence rate, X is a vector of explanatory variables implied by the model, and γ is a vector of parameters that are simple functions of the underlying technology parameters α and β.

One of the maintained assumptions of MRW is that g is constant across countries. Given the Cobb-Douglas production technology, TFP growth is equal to g times the exponent on the efficiency index, which here is $1-\alpha-\beta$. In the presence of wage differentials, TFP growth will be a function of structural change terms, so our extension of MRW takes the form:

$$\begin{aligned}\log\frac{Y(t)}{L(t)}-\log\frac{Y(0)}{L(0)} &= \omega+\frac{t\,(k-1)\,\phi}{1-\alpha-\beta}MGROWTH \\ &\quad +\frac{tk\phi}{(1-\alpha-\beta)\,\psi}DISEQ+\theta\gamma' X-\theta\log\frac{Y(0)}{L(0)}\end{aligned} \tag{22}$$

We use this specification in much of the empirical work. It provides a useful way to estimate growth in aggregate efficiency without using capital stock data. That said, its linear form relies on the Cobb-Douglas production function, the one-sector structure of the model, and the simple steady-state solution to which it gives rise. The specification (22) is therefore a hybrid of the Solow model and a two-sector framework of the kind set out in the main text. Although not wholly satisfactory, this reflects a long-standing difficulty in deriving a two-sector growth model that is simple enough to implement empirically.

The other necessary approximations are less serious. One of the explanatory variables in the MRW growth regression is $\log(n+g+\delta)$ where n is population or labor force growth, δ is depreciation and $g+\delta$ is typically assumed to equal 0.05. Our model, in which g varies across countries, weakens the case for treating $g+\delta$ in this way. In principle, a solution would be to substitute $MGROWTH$ and $DISEQ$ into the $\log(n+g+\delta)$ term and estimate the model by nonlinear least squares, but this model would be only weakly identified. An alternative and more pragmatic response is to argue that variation in g is likely to be modest in relation to the international variation in population growth rates (n).

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Table 1

Employment and output shares of agriculture in 1960, 1980 and 1996

	a			*s*			*RLP*			Sample sizes		
	1960	1980	1996	1960	1980	1996	1960	1980	1996	*a*	*s*	*RLP*
Sub-Saharan Africa	0.88	0.76	0.71	0.39	0.30	0.37	11.8	8.7	6.1	19	16	16
Middle East and North Africa	0.55	0.28	0.19	–	–	–	–	–	–	4	0	0
East Asia and Pacific	0.62	0.39	0.17	0.29	0.19	0.08	3.4	3.3	2.8	10	8	7
South Asia	0.75	0.70	0.60	0.46	0.34	0.25	3.2	3.2	3.4	5	4	4
Latin America and the Caribbean	0.53	0.36	0.23	0.23	0.12	0.09	3.8	3.4	2.2	20	18	18
High income OECD	0.19	0.08	0.05	0.11	0.05	0.03	2.4	1.8	1.7	20	16	16

Notes. Medians within each country grouping. Own calculations based on FAO (2003) and World Bank (2002); see text for details. a = share of agriculture in total employment. s = share of agriculture in total value added. RLP = ratio of average labor productivity in non-agriculture to that in agriculture (cf. text). Sample sizes = Number of countries in the regional sample.

Table 2

Propensity to migrate, 1960-96

Five highest	
Singapore	9.8%
Japan	5.2%
France	4.7%
Canada	4.4%
Korea, Republic of	4.4%
Five lowest	
Malawi	0.29%
Ghana	0.25%
Mozambique	0.22%
Niger	0.17%
Nepal	0.05%

Notes. Calculated as -(log(a96)-log(a60))/36 where aYY is the agricultural employment share in year 19YY.

Table 3

Sample correlations

	DY	ln(*Inv*)	ln(*YRSCH*)	ln(n+g+δ)	ln(*GDP60*)	*MGROWTH*	*DISEQ*	*MGROWTH2*	*DISEQ2*	*a60*
DY	1.00									
ln(*Investment)*	0.58	1.00								
ln(*YRSCH*)	0.41	0.71	1.00							
ln(n+g+δ)	-0.18	-0.23	-0.29	1.00						
ln(*GDP60*)	0.06	0.50	0.76	-0.39	1.00					
MGROWTH	0.26	0.22	0.18	0.25	0.16	1.00				
DISEQ	0.42	0.46	0.40	-0.01	0.40	0.81	1.00			
MGROWTH2	0.35	0.16	0.19	0.02	0.16	0.81	0.64	1.00		
DISEQ2	0.50	0.48	0.45	-0.20	0.44	0.73	0.93	0.76	1.00	
a60	-0.24	-0.59	-0.84	0.44	-0.90	0.00	-0.34	-0.04	-0.39	1.00

Notes. Correlations are for 76-country sample except for MGROWTH2 and DISEQ2 (66 countries). DY is the log difference of GDP per worker, 1960-96. ln(Investment) is the log of the investment share of GDP per capita, averaged over 1960-96. ln(YRSCH) is the log of the average years of schooling in the working-age population, averaged over 1960-95. ln(n + g + δ) includes labor-force growth (n), productivity growth (g) and depreciation (δ), 1960-96, where g + δ = 0.05 as in MRW (1992). ln(GDP60) is the log of GDP per worker in 1960. MGROWTH, DISEQ, MGROWTH2, and DISEQ2 are structural change terms calculated for 1960-96, as defined in the text. a60 is the agricultural employment share in 1960.

Table 4 – Structural change effects in the MRW model

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Period	1960-85	1960-85	1960-85	1960-85	1960-96	1960-96	1960-96	1960-80	1980-96
Observations	92	76	76	76	76	76	76	76	76
ln(Investment)	0.36*	0.34*	0.33*	0.25*	0.39*	0.47*	0.40*	0.25*	0.35+
	(0.06)	*(0.07)*	*(0.08)*	*(0.07)*	*(0.14)*	*(0.16)*	*(0.14)*	*(0.05)*	*(0.13)*
ln(SCHOOL)	0.25*	0.30*							
	(0.06)	*(0.07)*							
ln(YRSCH)			0.24*	0.26*	0.41*	0.39*	0.41*	0.15*	0.16
			(0.09)	*(0.07)*	*(0.10)*	*(0.11)*	*(0.10)*	*(0.05)*	*(0.11)*
ln(n+g+d)	-0.43º	-0.54+	-0.41	-0.70+	-0.97+	-1.14*	-1.00*	-0.49º	-0.75*
	(0.24)	*(0.24)*	*(0.26)*	*(0.30)*	*(0.38)*	*(0.38)*	*(0.29)*	*(0.25)*	*(0.21)*
ln(Initial GDP)	-0.29*	-0.29*	-0.25*	-0.32*	-0.44*	-0.40*	-0.43*	-0.20*	-0.22+
	(0.06)	*(0.06)*	*(0.08)*	*(0.07)*	*(0.10)*	*(0.10)*	*(0.10)*	*(0.05)*	*(0.08)*
MGROWTH				0.28	-0.08	0.67+		0.44º	-0.02
				(0.32)	*(0.53)*	*(0.26)*		*(0.23)*	*(0.24)*
DISEQ				1.14	2.14º		1.97*	0.28	0.45
				(0.69)	*(1.13)*		*(0.42)*	*(0.46)*	*(0.41)*
Constant	3.33*	3.14*	1.88+	1.31	1.86	1.19	1.76	1.02	0.54
	(0.73)	*(0.78)*	*(0.87)*	*(1.01)*	*(1.50)*	*(1.50)*	*(1.18)*	*(0.84)*	*(1.01)*
R^2	0.51	0.50	0.43	0.59	0.56	0.54	0.56	0.60	0.40
s.e.	0.31	0.30	0.32	0.28	0.40	0.41	0.40	0.23	0.27
F-test:				28.20	10.71			26.91	1.99
Prob.>F				*0.0000*	*0.0001*			*0.0000*	*0.1442*

*Notes. Dependent variable: log difference of GDP per worker over the specified period. White heteroscedasticity-consistent standard errors in parentheses. MGROWTH re-scaled by multiplying by 100; DISEQ re-scaled by multiplying by 1000. Significance level: * 1%; + 5%; ° 10%.*

Table 5

Structural change effects in the MRW model, 1960-96: further evidence

	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)
Observations	76	76	76	76	66	66	66	56	56	56
ln(Investment)	0.28º	0.33+	0.28º	0.28º	0.33º	0.39+	0.31º	0.29º	0.33º	0.30º
	(0.15)	*(0.16)*	*(0.15)*	*(0.15)*	*(0.18)*	*(0.19)*	*(0.17)*	*(0.16)*	*(0.17)*	*(0.16)*
ln(YRSCH)	0.39*	0.37*	0.39*	0.26º	0.40*	0.40*	0.40*	0.40*	0.39*	0.39*
	(0.10)	*(0.11)*	*(0.10)*	*(0.15)*	*(0.13)*	*(0.14)*	*(0.13)*	*(0.11)*	*(0.12)*	*(0.11)*
ln(n+g+d)	-1.20*	-1.18*	-1.22*	-1.10*	-0.88+	-0.83+	-0.89+	-1.30º	-0.96	-1.35+
	(0.39)	*(0.39)*	*(0.37)*	*(0.39)*	*(0.37)*	*(0.40)*	*(0.37)*	*(0.67)*	*(0.72)*	*(0.64)*
ln(GDP 1960)	-0.37*	-0.32*	-0.36*	-0.50*	-0.33*	-0.29*	-0.35*	-0.36*	-0.29+	-0.36*
	(0.10)	*(0.10)*	*(0.09)*	*(0.13)*	*(0.10)*	*(0.10)*	*(0.10)*	*(0.11)*	*(0.11)*	*(0.11)*
MGROWTH	-0.16	0.62*		0.41				-0.23	0.44	
	(0.45)	*(0.21)*		*(0.51)*				*(0.53)*	*(0.38)*	
DISEQ	2.08+		1.76*	1.11				2.26º		1.84+
	(0.99)		*(0.39)*	*(1.01)*				*(1.29)*		*(0.77)*
MGROWTH2					0.38	1.48*				
					(0.90)	*(0.44)*				
DISEQ2					3.31		4.09*			
					(2.19)		*(1.06)*			
a60				-0.86						
				(0.68)						
Constant	0.81	0.47	0.68	2.70	1.26	1.03	1.38	0.52	0.78	0.30
	(1.48)	*(1.45)*	*(1.36)*	*(1.87)*	*(1.41)*	*(1.43)*	*(1.34)*	*(1.91)*	*(1.98)*	*(1.74)*
R-squared	0.66	0.64	0.66	0.67	0.69	0.67	0.69	0.62	0.60	0.62
s.e.	0.37	0.38	0.37	0.37	0.38	0.38	0.37	0.42	0.43	0.42
F-test:	9.98			12.06	7.94			2.80		
Prob.>F	*0.0002*			*0.0000*	*0.0009*			*0.0713*		

Notes. Dependent variable: log difference of GDP per worker, 1960-96. All regressions include regional dummies (coefficients not reported). White heteroscedasticity-consistent standard errors in parentheses. MGROWTH and MGROWTH2 re-scaled by multiplying by 100; DISEQ and DISEQ2 re-scaled by multiplying by 1000. Significance level: * *1%;* $^{+}$ *5%; ° 10%.*

Table 6

Structural change effects on TFP growth

	(20)	(21)	(22)	(23)	(24)	(25)	(26)	(27)	(28)	(29)
TFP series	KR	KR	KR	KR	KR	KR	BG	BG	BG	BC
Period	1960-85	1960-85	1960-85	1960-85	1960-85	1960-85	1965-95	1965-95	1965-95	1960-2000
Observations	75	75	75	66	66	66	75	48	48	61
MGROWTH	0.48	1.04*					-1.62			
	(0.58)	*(0.21)*					*(1.09)*			
DISEQ	1.54		2.51*				5.46+			
	(1.27)		*(0.42)*				*(2.25)*			
MGROWTH2				-0.54	1.40*			-1.24		-0.23
				(0.73)	*(0.51)*			*(2.21)*		*(1.27)*
DISEQ2				6.71*		5.53*		7.45	5.20°	4.88°
				(1.90)		*(1.20)*		*(4.86)*	*(2.74)*	*(2.56)*
ln(*TFP60*)	-1.16*	-1.04*	-1.21*	-1.31*	-1.05*	-1.26*	-1.43*	-1.48*	-1.47*	
	(0.30)	*(0.30)*	*(0.31)*	*(0.29)*	*(0.30)*	*(0.27)*	*(0.38)*	*(0.48)*	*(0.47)*	
Constant	-0.07	-0.11	0.03	-0.04	-0.13	-0.10	-0.41	-0.90	-1.08*	0.86*
	(0.22)	*(0.21)*	*(0.19)*	*(0.20)*	*(0.22)*	*(0.17)*	*(0.51)*	*(0.54)*	*(0.40)*	*(0.25)*
R^2	0.50	0.49	0.49	0.54	0.47	0.54	0.47	0.36	0.35	0.46
s.e.	*0.46*	*0.46*	*0.46*	*0.45*	*0.47*	*0.44*	*0.94*	*1.14*	*1.13*	*0.58*
F(stru. change)	19.61			11.11			3.60	1.85		3.60
Prob. > *F*	*0.0000*			*0.0001*			*0.0326*	*0.1706*		*0.0342*

*Notes. Dependent variable: Average annual growth rate of total factor productivity, in percent. All regressions control for regional dummies; coefficients on regional dummies not reported. White heteroscedasticity-consistent standard errors in parentheses. MGROWTH and MGROWTH2 re-scaled by multiplying by 100; DISEQ and DISEQ2 re-scaled by multiplying by 1000. TFP series: KR = Klenow and Rodriguez-Clare (1997); BG = Bernanke and Gürkaynak (2001); BC = Bosworth and Collins (2003). Significance level: * 1%; + 5%; ° 10%.*

Table 7

Structural change and growth: instrumental variable results (Model 2)

Regression	(30)	(31)	(32)	(33)
Time period	1960-96	1960-96	1980-96	1980-96
Observations	66	66	61	61
MGROWTH2				
OLS	1.58*		0.38º	
	(0.44)		*(0.21)*	
2SLS	2.78+		0.49	
	(1.24)		*(0.29)*	
GMM	2.77+		0.47	
	(1.16)		*(0.29)*	
Fuller (1)	2.64+		0.49	
	(1.16)		*(0.35)*	
DISEQ2				
OLS		4.09*		0.83º
		(1.06)		*(0.46)*
2SLS		9.82+		1.45º
		(4.03)		*(0.86)*
GMM		9.47+		1.46º
		(3.97)		*(0.86)*
Fuller (1)		8.84+		1.41
		(4.28)		*(1.04)*
Robust 95% confidence interval	(-0.02, 7.76)	(n/a)	(-0.23, 1.27)	(-0.96, 4.47)
Instrument set	POPAGE LRLP60	POPAGE LRLP60	POPAGE MG26080	POPAGE DISEQ26080
First stage F-statistic	4.40	3.39	23.37	7.83
Sargan P-value	0.98	0.63	0.50	0.63
J-statistic P-value	0.98	0.64	0.39	0.55
Pagan-Hall P-value	0.11	0.33	0.04	0.05
Wu-Hausman P-value	0.26	0.19	0.79	0.56

Notes. Dependent variable: log difference of GDP per worker. Entries in table are coefficients and standard errors on either (rescaled) MGROWTH2 or DISEQ2 in growth regressions, treating that term as endogenous. Coefficients are not directly comparable across the two different time periods. All regressions include regional dummies and the MRW controls; these coefficients not reported. Significance level: * 1%; $^{+}$ 5%; $^{\circ}$ 10%. Fuller (1) is the Fuller modification of LIML with alpha=1 (see text). Robust confidence intervals constructed using conditional likelihood ratio tests developed by Moreira (2003). Null hypotheses of specification tests are overidentifying restrictions valid (Sargan for 2SLS, J-statistic for GMM); system homoskedastic (Pagan-Hall, for 2SLS); regressor exogenous (Wu-Hausman, for 2SLS).

Table 8

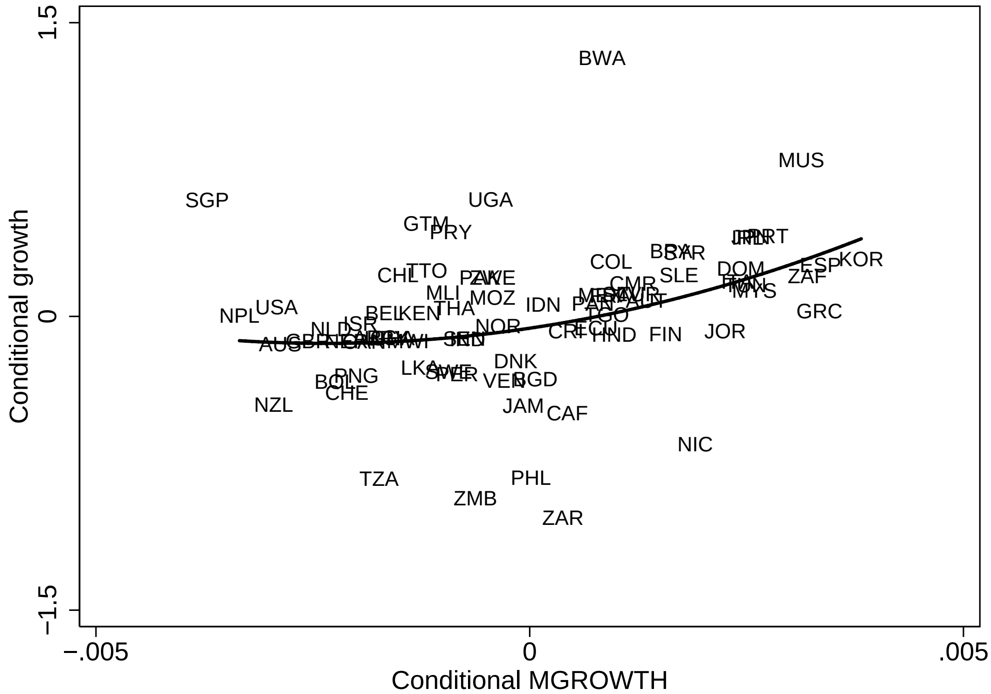
Implied parameter values in the growth and TFP regressions

Dependent variable	Growth		TFP-KR		TFP-BG		Growth		TFP-BG	
Time period	1960-96		1960-85		1965-95		60-80	80-96	65-80	80-95
Sample	all	dev.	all	dev.	all	dev.	all	all	all	all
Restricted model with $\psi = \infty$										
k	1.87	1.61	2.56	2.68	1.79	1.76	2.31	1.51	2.88	1.17
Restricted model with $k = 1$										
ψ	0.041	0.039	0.027	0.022	0.028	0.023	0.038	0.080	0.024	0.044
w_m/w_a										
10th percentile	1.08	1.08	1.13	1.14	1.12	1.15	1.07	1.05	1.13	1.09
Median	1.50	1.32	1.76	1.60	1.78	1.55	1.51	1.22	1.86	1.40
90th percentile	2.06	1.90	2.64	2.60	2.61	2.64	2.22	1.55	3.00	1.99

Notes. Model 1 specification. All underlying growth regressions control for the four MRW variables and regional dummies. All underlying TFP regressions control for regional dummies and initial TFP.

Figure 1

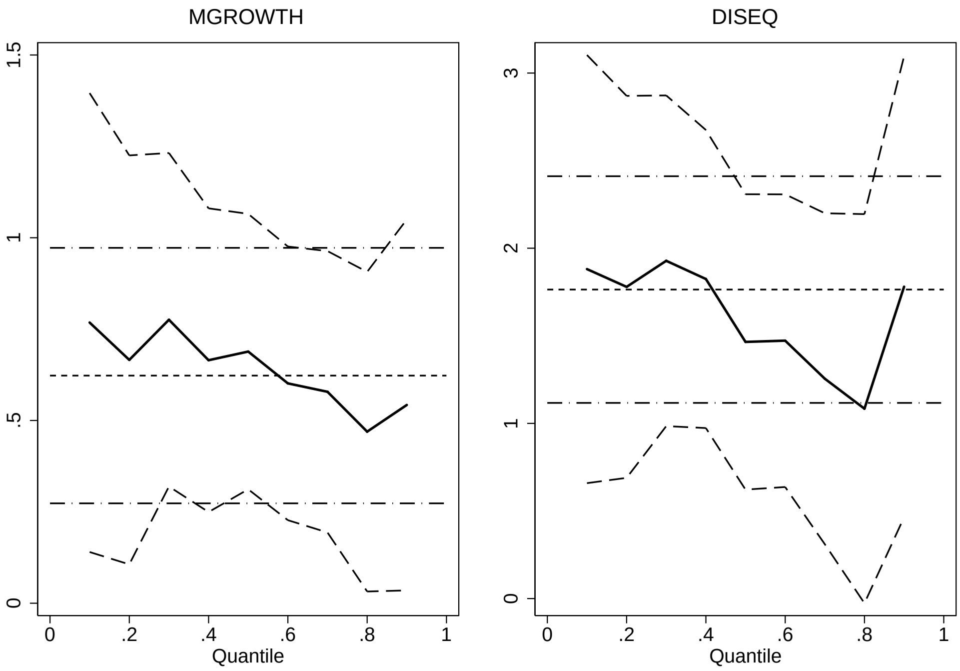
Convexity in the relationship between growth and structural change, 1960-96



Notes. Vertical axis: DY, conditional on the four MRW variables and regional dummies. Horizontal axis: MGROWTH, conditional on the same variables. Sample: 76 countries. Quadratic regression line added, ignoring the three outliers BWA, SGP and ZAR.

Figure 2

Quantile regression estimates for the structural change terms



Notes. The underlying models correspond to regressions (11) and (12) of Table 5, respectively. The solid curve represents the coefficient estimates for each decile, and the two dashed lines represent the 90 percent confidence bands for these estimates. The straight dashed line shows the OLS estimate, with its 90 percent confidence interval represented by the two straight dash-dotted lines.

Appendix

Table A1

Structural change effects in the MRW model, 1960-96: median and robust regression results

	(A1)	(A2)	(A3)	(A4)	(A5)	(A6)	(A7)	(A8)	(A9)
Estimation	MR	MR	MR	MR	MR	MR	RR	RR	RR
Observations	76	76	76	66	66	66	76	66	65
ln(Investment)	0.22*	0.29*	0.22	0.31	0.42*	0.23+	0.23*	0.29*	0.28*
	(0.08)	*(0.09)*	*(0.14)*	*(0.22)*	*(0.07)*	*(0.11)*	*(0.09)*	*(0.10)*	*(0.09)*
ln(YRSCH)	0.39*	0.39*	0.40*	0.36	0.29*	0.39*	0.41*	0.42*	0.44*
	(0.08)	*(0.10)*	*(0.15)*	*(0.24)*	*(0.08)*	*(0.11)*	*(0.09)*	*(0.10)*	*(0.10)*
ln(n+g+d)	-1.13*	-1.30*	-1.11°	-1.22	-1.17*	-1.10+	-1.08*	-0.79°	-0.85+
	(0.38)	*(0.42)*	*(0.60)*	*(0.86)*	*(0.30)*	*(0.51)*	*(0.39)*	*(0.44)*	*(0.42)*
ln(GDP 1960)	-0.34*	-0.35*	-0.37*	-0.34+	-0.28*	-0.36*	-0.40*	-0.38*	-0.38*
	(0.06)	*(0.07)*	*(0.11)*	*(0.16)*	*(0.05)*	*(0.08)*	*(0.07)*	*(0.07)*	*(0.07)*
MGROWTH	0.24	0.69*					0.19		
	(0.33)	*(0.19)*					*(0.34)*		
DISEQ	1.09		1.47+				1.19		
	(0.77)		*(0.62)*				*(0.77)*		
MGROWTH2				0.57	1.29*			0.25	0.83
				(1.44)	*(0.26)*			*(0.66)*	*(0.72)*
DISEQ2				2.20		3.28*		2.79°	1.77
				(3.64)		*(1.12)*		*(1.65)*	*(1.74)*
Constant	0.60	0.25	1.01	0.43	0.30	0.81	1.18	1.80	1.48
	(1.12)	*(1.26)*	*(1.80)*	*(2.64)*	*(0.90)*	*(1.48)*	*(1.18)*	*(1.27)*	*(1.24)*
F-test:	8.42			1.44			7.76	6.49	8.36
Prob.>F	0.0006			0.2462			0.0009	0.0030	0.0007

Notes. Dependent variable: log difference of GDP per worker, 1960-96. MR is median regression, RR an outlier-robust estimator. All regressions include regional dummies (coefficients not reported). White heteroscedasticity-consistent standard errors in parentheses. MGROWTH and MGROWTH2 re-scaled by multiplying by 100; DISEQ and DISEQ2 re-scaled by multiplying by 1000. Regression (A9) excludes Singapore. Significance level: * 1%; $^{+}$ 5%; ° 10%.

Table A2

Structural change effects in developing countries, 1960-96

	(A10)	(A11)	(A12)	(A13)	(A14)	(A15)	(A16)
Estimation	OLS	OLS	OLS	OLS	MR	MR	MR
Observations	48	48	48	56	56	56	56
ln(Investment)	0.34	0.39°	0.32°	0.26	0.22°	0.31+	0.33*
	(0.21)	*(0.22)*	*(0.19)*	*(0.16)*	*(0.12)*	*(0.14)*	*(0.12)*
ln(YRSCH)	0.40+	0.41+	0.40*	0.33*	0.39*	0.36+	0.27+
	(0.15)	*(0.16)*	*(0.14)*	*(0.12)*	*(0.12)*	*(0.15)*	*(0.13)*
ln(n+g+d)	-0.97	-0.67	-0.99	-0.67	-1.13	-1.47°	-1.01
	(0.64)	*(0.66)*	*(0.64)*	*(0.67)*	*(0.73)*	*(0.87)*	*(0.66)*
ln(Initial GDP)	-0.33*	-0.26+	-0.34*	-0.46*	-0.34*	-0.35*	-0.30*
	(0.12)	*(0.11)*	*(0.11)*	*(0.12)*	*(0.09)*	*(0.11)*	*(0.10)*
MGROWTH				-0.70	0.24	0.78°	
				(0.56)	*(0.49)*	*(0.41)*	
DISEQ				2.36°	1.09		1.55°
				(1.20)	*(1.28)*		*(0.80)*
MGROWTH2	0.39	1.41+					
	(1.06)	*(0.63)*					
DISEQ2	3.62		4.46+				
	(3.01)		*(1.80)*				
*MGROWTH*SSAfrica*				1.05			
				(1.87)			
*DISEQ*SSAfrica*				2.79			
				(4.54)			
Constant	1.03	1.22	1.09	3.19	0.60	-0.19	0.99
	(1.78)	*(1.83)*	*(1.73)*	*(2.10)*	*(1.89)*	*(2.26)*	*(1.63)*
R-squared	0.66	0.64	0.66	0.67			
s.e.	0.44	0.44	0.43	0.41			
F-test:	3.40			2.14	2.03		
Prob.>F	*0.0439*			*0.0922*	*0.1426*		

Notes. Dependent variable: log difference of GDP per worker, 1960-96. All regressions include regional dummies (coefficients not reported). White heteroscedasticity-consistent standard errors in parentheses. MGROWTH and MGROWTH2 re-scaled by multiplying by 100; DISEQ and DISEQ2 re-scaled by multiplying by 1000. Significance level: * *1%;* $^{+}$ *5%; ° 10%.*

Table A3

Structural change effects on TFP growth: sub-periods and developing countries

	(A17)	(A18)	(A19)	(A20)	(A21)	(A22)	(A23)	(A24)	(A25)
TFP series	BG	BG	BG	BG	KR	KR	BG	BG	BC
Period	1965-80	1965-80	1965-80	1980-95	1960-85	1960-85	1965-95	1965-95	1960-2000
Observations	75	75	75	75	55	48	66	41	43
MGROWTH	0.35	1.25*		-1.93+	0.48		-2.38+		
	(0.99)	*(0.43)*		*(0.96)*	*(0.72)*		*(1.00)*		
DISEQ	2.19		2.79*	4.67+	2.02		6.56*		
	(1.83)		*(0.74)*	*(1.84)*	*(1.77)*		*(2.09)*		
MGROWTH2						-0.86		-2.40	-0.95
						(0.84)		*(2.00)*	*(1.36)*
DISEQ2						9.50*		8.60°	5.64°
						(3.20)		*(4.78)*	*(2.89)*
ln(*TFP60*)	-1.62*	-1.51*	-1.65*	-0.68	-1.29*	-1.50*	-1.23*	-1.31°	
	(0.43)	*(0.42)*	*(0.42)*	*(0.57)*	*(0.44)*	*(0.39)*	*(0.44)*	*(0.60)*	
Constant	-1.56*	-1.64*	-1.47*	0.60	-0.19	-0.17	0.25	-0.18	1.00*
	(0.55)	*(0.54)*	*(0.51)*	*(0.75)*	*(0.33)*	*(0.28)*	*(0.39)*	*(0.57)*	*(0.28)*
R^2	0.40	0.38	0.40	0.37	0.44	0.48	0.51	0.37	0.32
s.e.	*1.12*	*1.13*	*1.12*	*1.26*	*0.53*	*0.51*	*0.93*	*1.18*	*0.66*
F(stru. change)	7.40			3.23	6.98	5.34	4.94	1.64	2.03
Prob. > *F*	*0.0013*			*0.0457*	*0.0022*	*0.0087*	*0.0104*	*0.2091*	*0.1452*

Notes. Dependent variable: Average annual growth rate of total factor productivity, in percent. All regressions include regional dummies (coefficients not reported). White heteroscedasticity-consistent standard errors in parentheses. MGROWTH and MGROWTH2 re-scaled by multiplying by 100; DISEQ and DISEQ2 re-scaled by multiplying by 1000. TFP series: KR = Klenow and Rodriguez-Clare (1997); BG = Bernanke and Gürkaynak (2001); BC = Bosworth and Collins (2003). Significance level: * *1%;* $^+$ *5%; ° 10%.*

Table A4

NLS regressions

Dependent variable	Growth		TFP-KR	
Sample	all	developing	all	developing
Restricted model with $\psi = \infty$				
	(A26)	(A27)	(A28)	(A29)
k	2.07*	1.66^{+}	4.63*	4.55^{+}
	(0.68)	*(0.77)*	*(1.58)*	*(1.84)*
Restricted model with $k = 1$				
	(A30)	(A31)	(A32)	(A33)
ψ	0.0240°	0.0297	0.0044^{+}	0.0041^{+}
	(0.0126)	*(0.0238)*	*(0.0017)*	*(0.0018)*
w_m/w_a				
10th percentile	1.14	1.10	1.76	1.75
median	1.84	1.42	5.56	4.21
90th percentile	2.81	2.19	10.77	9.53

Notes. Dependent variable: average annual growth of GDP per worker, 1960-96, in the case of the growth regressions; average annual growth rate of total factor productivity, 1960-85, in the case of the TFP regressions. All regressions include regional dummies. The growth regressions additionally control for the four MRW variables (investment, schooling, ln(n+g+δ), and initial GDP). The TFP regressions additionally control for the log of initial TFP. Coefficients on control variables not reported. Significance level: * 1%; $^{+}$ 5%; ° 10%

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