Econometric Studies of Growth, Convergence and Conflicts

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Thesis submitted for the degree of Doctor of Philosophy,
Sub-Faculty of Economics, University of Oxford, Trinity Term 1998

Carpe Diem

To Jill in loving memory

Abstract

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This thesis presents a methodological examination of growth regressions and an econometric analysis of civil wars.

Commonly used single cross-country growth regressions are likely to suffer from omitted variable and endogeneity bias. Using panel data we estimate the (augmented) Solow growth model. Our results suggest that unobserved country specific effects are significant and that some of the explanatory variables, in particular investment, should not be treated as strictly exogenous. Based on a comparison of different estimation methods we suggest the use of a recently developed system generalized method of moments (GMM) estimator for cross-country growth regressions. Using this preferred method of estimation we show that the estimated speed of convergence is higher than suggested by OLS levels estimation, but lower than the estimates obtained from first differenced GMM estimation.

We use this preferred method of estimation to examine two further empirical growth issues. First, we present a study of openness, investment and growth and second, we examine the relationship between growth, geographical location, and some cultural characteristics. Our analysis suggests that once we account for endogeneity and unobserved country specific effects we can only confirm some but not all of the commonly found results in the literature. For example, we find that using our preferred estimation method the augmented Solow model can account for the low growth performance of Sub-Saharan African and Latin American countries.

In addition, we present an economic analysis of civil wars. We construct a simple microeconomic model of the causes of civil wars and test our hypotheses using a global data set. The duration of civil wars and post-war peace is examined by using proportional hazard models. We find that both the causes and the duration of civil wars are related to socio-economic variables, although these two processes are distinct.

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Chapter 1

Introduction

1.1 Research Objectives

This thesis considers specific estimation problems in the econometric analysis of growth and presents an econometric analysis of civil wars.

The main contribution of the thesis within the analysis of economic growth is a methodological examination of cross-country growth regressions. In the existing literature single cross-section growth regressions, estimated by ordinary least squares (OLS), are the most commonly used approach to studying the empirics of economic growth. We suggest that single cross-section growth regressions suffer both from omitted variable and endogeneity bias and thus may provide misleading results with regard to the estimated speed of convergence and the significance of some of the explanatory variables. Based on the framework of the (augmented) Solow growth model we examine these sources of bias by comparing alternative estimation methods. Our results suggest that unobserved country specific effects are significant and that some of the explanatory variables, in particular investment, should not be treated as strictly exogenous. We suggest the use of a recently developed system generalized method of moments (GMM)

estimator (Blundell and Bond (1998a)) to address the problems of unobserved country specific effects and endogeneity. We provide evidence that failing to account for unobserved country specific effects and the endogeneity of some of the explanatory variables may cause the speed of convergence to be underestimated. However, we also show that the recently suggested use of the Arellano and Bond (1991) first differenced GMM estimator for growth regressions provides estimated speeds of convergence which are likely to be too high. System GMM estimation produces more precise and more reasonable coefficient estimates than the differenced GMM estimator. Thus, we suggest the use of the system GMM estimator as our preferred method of estimation.

The discussion of these methodological issues and their implications for the estimated speed of convergence is followed by the examination of two further growth issues. Our objective is to examine whether commonly obtained results in the literature are confirmed when we use our preferred method of estimation. First, we examine the relationship between openness, investment and growth. Single cross-country growth regressions confirm the claim by Sachs and Warner (1997) that the coefficient on investment is not significant when the model controls for openness to international trade. We question this finding, because using system GMM estimation we find that once we account for endogeneity and unobserved country specific effects there are strong positive effects of both investment and openness on income growth. Second, we consider the relationship between growth, geographical location, and some cultural characteristics. In our study we confirm some but not all of the conventional results. For example, we find that using system GMM estimation the augmented Solow model can account for the low growth performance of Sub-Saharan African and Latin American countries. In summary, this thesis provides evidence that some of the earlier results in the literature may be seriously

biased due to unobserved country specific effects and endogeneity.

We then turn to an econometric examination of civil wars. The existing literature on civil wars falls into two distinct parts. On the one hand there are a variety of rational choice models explaining individual's participation in civil wars. On the other hand there is a large number of case studies. The main contribution of this thesis is to create a simple microeconomic framework which can be used for an empirical examination of the causes of civil wars, the duration of civil wars and the duration of post-war peace. To our knowledge these issues have so far not been investigated econometrically. Using a recently developed global data set of civil wars we examine the causes of civil wars in a probit model. In addition we use proportional hazard models to examine the duration of civil wars and post-war peace. We find that both the causes and the duration of civil wars are related to socio-economic variables, although these two processes are distinct.

In the remainder of this introductory chapter we provide an overview of the data sets used and a summary of the main research findings for each chapter.

1.2 Data Sources

Alongside the burgeoning growth literature a large number of cross country data sets have been compiled many of which are available through the internet. The most commonly used data source is the Penn World Table (PWT).[1] It offers annual GDP and components of GDP data for 152 countries between 1960 and 1992. All expenditure entries are denominated in a common set of prices in a common currency so that real international quantity comparisons can be made

[1] For a detailed discussion of PWT data please refer to Summers and Heston (1991, 1996) and the literature cited there.

both between countries and over time. Instead of using the relative prices of one country, the world structure of prices is considered with the aim of accounting for differences in purchasing power parity across countries. The GDP data in PWT is based on the set of benchmark studies of the United Nations International Comparison Program. The most recent version of PWT is Mark 5.6 which was used throughout this study.

The World Bank also provides cross-country time series data. The World Development Indicators (WDI) are available on CD-Rom and provide economic, financial and socio-economic data for most countries from 1970 to 1996 (WDI 1997).

Furthermore, a number of researchers have compiled their own data sets which are often publicly available. They are too numerous to list here, but we would like to thank Jeffrey Sachs and Andrew Warner as well as Robert Barro for making their data sets available to us.

For our economic analysis of civil wars we use the comprehensive Singer and Small (1994) data set. This global data set provides information about wars fought between 1816 and 1992.

We discuss the choice of data and data sources used in detail in each chapter.

1.3 Summary of the Main Results

1.3.1 Chapter 2

Chapter 2 constitutes an introductory chapter for Chapters 3, 4 and 5 in which we consider a number of economic growth issues. In Chapter 2 we present a selective survey of the empirical growth literature and we concentrate on the models and issues which we want to examine in the subsequent chapters. A common characteristic of empirical growth studies is that the empirics are not preceded by detailed theoretical modelling. One important exception is the discussion and testing of the Solow growth model. Most empirical modelling is more *ad hoc* and detailed

testing of the propositions of the so called 'new growth theory' is noticeably absent.

In the literature there is strong empirical evidence for conditional convergence, i.e. controlling for other variables the coefficient on initial income in a growth equation is significant and negative. This is interpreted as a 'catch up' mechanism, *ceteris paribus* poorer countries will tend to catch up with richer countries. There is also strong evidence that investment in physical capital and openness to international trade are positive for growth. However, the evidence for a positive effect from human capital and institutions is relatively weak. It is difficult to find appropriate proxies for human capital and institutions and the fragility of the results may be due to conceptual problems and data quality issues.

The most popular estimation method used in the empirical growth literature is cross-section regression, using average values for 1960-85 or 1960-90 for each country. These cross-country regressions can neither account for the presence of unobserved country specific effects nor for the endogeneity of (some of) the regressors. The panel data study by Caselli, Esquivel and Lefort (1996) finds that not accounting for these problems introduces an upward bias on the coefficient on initial income. As a result, single cross-country regressions tend to underestimate the speed of convergence.

1.3.2 Chapter 3

In Chapter 3 we turn to the methodological discussion of cross-country growth regressions. We base our discussion on empirical tests of the Solow model as well as its augmented version (which includes a measure of human capital as an additional input in the production function). We chose these models, because they have been examined by a number of researchers, most notably Mankiw, Romer and Weil (1992), Islam (1995) and Caselli, Esquivel and Lefort

(1996). Choosing an already well discussed model seemed appropriate for our methodological examination. If we had chosen to create our own growth model it would have been much more difficult to illustrate our concerns regarding estimation methods, since any differences in results could then also be attributed to the difference in models and not only to differences in the methodology.

We use a panel data set with repeated observations on five year periods and show that OLS levels estimation is likely to suffer from biases due to simultaneity and the presence of unobserved country specific effects. A within groups estimator will also produce biased and inconsistent estimates. Furthermore, we highlight a potential problem when using Caselli, Esquivel and Lefort's (1996) preferred first differenced GMM estimator. Blundell and Bond (1998a) provide evidence based on Monte Carlo studies which shows that when the series are persistent and the number of time series observations is small this first differenced GMM estimator can be poorly behaved, since past levels of the series provide weak instruments for subsequent first differences. We find that the estimated coefficient on initial income is very close to that obtained using the within groups estimator, suggesting that the first differenced GMM estimate suffers from a considerable downward bias. As a result, the estimated speed of convergence based on the first differenced GMM estimates is likely to be too high. As an alternative we consider the use of the Blundell and Bond (1998a) system GMM estimator which exploits stationarity properties to obtain more informative instruments. The estimated speed of convergence based on system GMM is much closer to the conventionally reported two percent *per annum*.

Moreover, we find investment in physical capital to be highly significant in all regressions, even though we control for the endogeneity of investment in the preferred system GMM results.

1.3.3 Chapter 4

Following the methodological discussion presented in Chapter 3 we then present an analysis of openness, investment and growth, building on the work by Sachs and Warner (1995, 1997). They claim that openness to international trade is such an important determinant of investment, that the coefficient on investment will be insignificant in a growth regression which already controls for openness. We focus on two possible explanations of why investment may not be significant in a single cross-country growth regression. First, openness and/or other regressors may be highly correlated with investment, thus, causing the estimated coefficient on investment to be insignificant. A second possible explanation is that ignoring unobserved country specific effects and endogeneity problems (which we found to be significant in the previous chapter) may cause investment to be insignificant in the Sachs and Warner model.

Different specifications of the single cross-country regressions show that it is the inclusion of life expectancy not openness which causes the coefficient on investment to be insignificant. However, growth regressions that include life expectancy as an explanatory variable are likely to suffer from endogeneity problems, and there may be omitted variables that affect both life expectancy and income. When we account for unobserved country specific effects and address the endogeneity issues using our preferred system GMM estimator, we find that investment is significant in the growth regressions despite controlling for openness. Thus, both our single cross-section and our panel regressions show that the Sachs-Warner measure of openness has a significant, positive effect on income. However, this variable is not so important that it drives investment out of the model.

1.3.4 Chapter 5

In chapter 5 we present a brief analysis of the relationship between growth and some measured country characteristics. We concentrate on truly time invariant characteristics such as geographical location, and cultural variables (religious affiliations and ethnic diversity) which only change very slowly over time.

First, we use the consistently estimated augmented Solow model coefficients from Chapter 3 to calculate the residuals of this growth regression. Second, we regress these residuals on the measured country specific characteristics. We use OLS to estimate this second regression. The estimated coefficients on the measured country characteristics are consistent if and only if all these measured country characteristics are uncorrelated with the unobserved country specific effects. Since this is a strong restriction, we only want to interpret these coefficient estimates as indicators of whether or not certain measured country characteristics are associated with high or low growth rates. We do not want to attach any causal significance to these results.

One of our findings is that the coefficients on dummy variables for African and Latin American countries are insignificantly different from zero. This suggests that the augmented Solow model can account for the growth performance of these countries once we allow for unobserved country specific effects in estimating the parameters of the model. A further interesting result is that countries with a high proportion of Eastern Religions experienced higher growth rates on average while countries with a high proportion of Protestants experienced lower growth rates on average. We do not find any evidence for a significant relationship between ethnic diversity and growth. Thus, we can confirm some but not all of the commonly found results regarding the relationship between measured country characteristics and growth.

1.3.5 Chapter 6

In the last chapter we present an econometric analysis of civil wars. First, we investigate whether civil wars have economic causes. We construct a simple model which is based on utility theory, in which rebels will conduct a civil war if the perceived benefits outweigh the costs of rebellion. Using a probit model we test our propositions empirically. We find that the coefficients on three variables, initial income, a measure of ethnic diversity and the initial population size are significant in this analysis of the probability of civil wars occurring.

Furthermore, we suggest that the same explanatory variables may also be related to the duration of civil war and postwar peace. We use proportional hazard models to test this hypothesis. Our results suggest that although the causes and the duration of civil war are related to socio-economic variables, these two processes are distinct. Income levels and population size are important causes of civil war, but have no effect on the duration of conflict and little effect on the duration of post-conflict peace. By contrast, ethnic diversity is both absolutely and relatively more important as a determinant of the duration of war and of postwar peace than as a cause of war. However, in all cases the effect of ethnic diversity is non-monotonic, both homogenous and highly diverse societies are substantially safer than those with middle range diversity. We explain this non-monotonic relationship in terms of the costs of coordination of rebellion, which are minimised when society is split between two ethnic groups. We explain the differences between the causes and duration of civil wars in terms of the irreversibility of the military recruitment decision and the problems posed by maintaining cohesion of the officer core during a conflict. The hazard of peace during a war falls sharply after the first year. We explain this in terms of the systematic over-optimism of actual rebels which is analogous to the winner's curse in auction theory.

Chapter 2

A Selective Survey of the Empirical Growth Literature

2.1 Introduction

The literature on the theory as well as the empirics of economic growth has increased enormously over the past decade. For instance the Social Sciences Citation Index lists 20,875 references for growth for the last 10 years. Thus, this literature survey can only be selective. We concentrate on discussion of models and variables which we examine in our empirical analysis, presented in Chapters 3, 4 and 5.

Within the theory of economic growth two main categories can be distinguished, the neo-classical growth models and the endogenous growth models. The neo-classical models are based on the work by Solow (1956) and there are a number of studies which test the propositions of this neo-classical growth theory. Since the neo-classical growth model assumes decreasing marginal returns to capital, one of the model's predictions is that initially poorer countries will tend to experience higher growth rates than richer countries: poorer countries will tend to 'catch up' and their income level will converge to the income levels of richer economies, *ceteris paribus*.

However, although empirical studies show that the Solow model fits several stylized facts, one limitation is that output per capita along the balanced growth path grows at the exogenously given rate of technological progress.

Clearly, the notion of exogenously given technological progress is unsatisfactory and Romer (1986) revived the growth debate by including the determination of technological progress in his growth model, i.e. technological progress was now endogenous to the model. Technological progress is modelled either by introducing research and development into the model or by assuming a learning-by-doing process. Depending on the other parameters of the model it is possible to have decreasing, constant, or increasing returns to production factors, i.e. the growth path can be stable or explosive. A further difference from the Solow model is that economic policy can have an impact on the long run rate of economic growth. (In the Solow model policy can increase savings, but this increase does not have a growth, only a level effect, in the long run.)

Parallel to this theoretical work exists a vast number of empirical studies, of which Barro (1991) 'Economic Growth in a Cross Section of Countries' is probably the best known. It is important to stress that Barro and most other authors do not explicitly set up a theoretical framework and test specific hypotheses developed in an endogenous growth model. Many of the empirical models have been constructed on an *ad hoc* basis. What Barro implicitly formulates in his work is a mixture of neo-classical and endogenous growth theory. For example the conditional convergence hypothesis is one of the main ideas derived from the neo-classical model. The inclusion of human capital and policy measures, however, give tribute to the ideas discussed in the endogenous growth literature.

Most of these 'Barro style' regressions are cross-country growth regressions, explaining av-

erage growth rates for 1960-85 or 1960-90 for a large cross-section of countries. The estimation method used is ordinary least squares (OLS), and a range of very different hypotheses has been tested. Growth studies analyse for example whether investment in physical and human capital, fiscal and monetary policy, openness to international trade, geographical location, democracy, ethnic diversity and religious affiliation can explain countries' growth performance. Levine and Renelt (1992) and Sala-i-Martin (1997a, 1997b) provide an overview of the different empirical results and examine their robustness. However, robustness of the results is not the only problematic issue with regard to growth regressions and a number of studies have recently suggested that the OLS estimation results of 'Barro style' growth regressions are likely to suffer from biases due to endogeneity and the presence of unobserved country specific effects.[1]

This review is structured in the following way. In the second section we discuss the predictions of the neo-classical growth model, in particular its predictions concerning income convergence. In the third section we discuss the determinants of growth as suggested by the '*ad hoc* empirical models'. In the fourth section we highlight some of the estimation problems. The last section concludes.

2.2 The Neo-Classical Growth Model and the Convergence Debate

In spite of the large number of endogenous growth models developed over the last ten years, the neo-classical growth model based on Solow (1956) remains a popular model in the debate about the empirics of economic growth. Solow's theoretical framework provides the basis for a

[1] For example Caselli, Esquivel and Lefort (1996) and Easterly, Loayza, and Montiel (1997) use a first differenced GMM estimator for panel data growth regressions in order to address the endogeneity and omitted variable biases.

number of empirical studies, such as Mankiw, Romer and Weil (1992) (henceforth MRW), Islam (1995) and Caselli, Esquivel and Lefort (1996) (henceforth CEL). In this section we present a detailed discussion of the Solow growth model.

2.2.1 The Textbook Solow Model

The Solow model is based on a Cobb-Douglas production function with three inputs; capital, K, labour, L and labour augmenting technological progress, A.[2] Assuming constant returns to scale we can write for continuous time

$$Y(t) = K(t)^{\alpha}(A(t)L(t))^{1-\alpha} \tag{2.1}$$

or in the intensive form

$$\widehat{y}(t) = \widehat{k}(t)^{\alpha} = f(\widehat{k}(t))$$

where

$$0 < \alpha < 1$$

and $\widehat{y}(t) = \frac{Y(t)}{A(t)L(t)}$ denotes output per effective unit of labour, and $\widehat{k}(t) = \frac{K(t)}{A(t)L(t)}$ denotes capital per effective unit of labour.

Capital is subject to diminishing marginal returns

$$\begin{aligned} f'(\widehat{k}(t)) &> 0 \\ f''(\widehat{k}(t)) &< 0 \end{aligned}$$

[2] The exposition of the textbook Solow model is mainly based on MRW, Romer (1996, chapter 1) and Barro and Sala-i-Martin (1995, chapter 1).

Labour force and technology grow at the following constant and exogenous rates

$$L(t) = L(0)e^{nt} \tag{2.2}$$

$$A(t) = A(0)e^{gt} \tag{2.3}$$

where n is the growth rate of the labour force and g is the rate of technological progress. We further assume that a constant fraction of output, s, is invested[3] and capital depreciates at a constant rate δ. Thus, the change of the capital stock over time can be written as

$$\dot{K}\ (t) = sY(t) - \delta K(t) \tag{2.4}$$

$$\text{where } \dot{K}\ (t) = \frac{dK(t)}{dt}.$$

From this we can derive how the capital stock per effective worker changes over time

$$\begin{aligned} \dot{\widehat{k}}\ (t) &= \frac{d\left(\frac{K(t)}{A(t)L(t)}\right)}{dt} = \frac{\dot{K}\ (t)A(t)L(t) - K(t)(A(t)\ \dot{L}\ (t) + L(t)\ \dot{A}\ (t))}{(A(t)L(t))^2} \\ &= \frac{\dot{K}\ (t)}{A(t)L(t)} - \frac{K(t)}{A(t)L(t)}\frac{\dot{L}\ (t)}{L(t)} - \frac{K(t)}{A(t)L(t)}\frac{\dot{A}\ (t)}{A(t)} \end{aligned}$$

Substituting from equation (2.4) for $\dot{K}\ (t)$ and substituting n for the growth rate of the work-force, $\frac{\dot{L(t)}}{L(t)}$, and g for the rate of technological progress, $\frac{\dot{A(t)}}{A(t)}$, this expression simplifies to

$$\dot{\widehat{k}}\ (t) = s\widehat{y}(t) - \widehat{k}(t)(n + g + \delta) \tag{2.5}$$

Equation (2.5) provides an equation of motion for the capital stock per effective unit of labour and it is the key equation of the Solow model. $s\widehat{y}(t)$ is the actual amount of investment per

[3]The Solow model describes the growth rate for a closed economy, thus investment equals savings.

effective unit of labour. $\widehat{k}(t)(n+g+\delta)$ is the investment per effective unit of labour which is necessary to keep the capital stock per effective unit of labour exactly at the previous level. Thus, we will refer to it as the break even investment per effective unit of labour. The difference between actual and break even investment per effective unit of labour determines the change in the capital stock per effective unit of labour. It is easy to show that these assumptions ensure that there is a steady state. When actual investment per effective unit of labour is higher than the investment needed to keep the capital stock per effective unit of labour at the previous level, the capital stock per effective unit of labour increases and $\dot{\widehat{k}}(t)$ is positive. However, since capital exhibits diminishing marginal returns the capital stock per effective unit of labour only increases up to the point when actual investment per effective unit of labour is equal to break even investment per effective unit of labour. If actual investment per effective unit of labour is lower than the break even investment per effective unit of labour, $\dot{\widehat{k}}(t)$ is negative, because savings are insufficient to maintain the previous capital stock. The capital stock per effective unit of labour will fall until actual and break even investment per effective unit of labour are equal. Thus, regardless at which capital stock per effective unit of labour the economy starts it will converge to the capital stock per effective unit of labour for which actual and break even investment per effective unit of labour are equal.

The capital stock per effective unit of labour which fulfills this requirement is the steady state capital stock per effective unit of labour, because $\dot{\widehat{k}}(t)$ is zero. In the case of a growing labour force and positive labour augmenting technological progress, the actual capital stock grows at rate $(n+g)$ in the steady state.

The assumption of diminishing marginal returns to capital ensures that there is a steady state and that no matter at which initial capital stock per effective unit of labour the economy

starts, it will converge to this steady state capital stock per effective unit of labour.

Formally, the steady state capital stock per effective unit of labour, $\widehat{k}^*$, can be determined by setting equation (2.5) equal to zero

$$\widehat{k}^* = \left(\frac{s}{(n+g+\delta)}\right)^{\frac{1}{1-\alpha}} \tag{2.6}$$

The steady state output per effective unit of labour can be derived by substituting (2.6) into the production function. Taking logarithms gives

$$\ln \widehat{y}^* = \left(\frac{\alpha}{1-\alpha}\right)(\ln s - \ln(n+g+\delta)) \tag{2.7}$$

The rate at which an economy reaches the steady state can be determined in the following way. Dividing both sides of equation (2.5) by $\widehat{k}(t)$ provides us with an expression for the growth rate of the capital stock per effective unit of labour, $\gamma_{\widehat{k}}$

$$\gamma_{\widehat{k}} = \frac{\dot{\widehat{k}}(t)}{\widehat{k}(t)} = \frac{sf(\widehat{k}(t))}{\widehat{k}(t)} - (n+g+\delta)$$

We can rewrite this equation as

$$\gamma_{\widehat{k}} = \frac{d\ln \widehat{k}(t)}{dt} = s \cdot e^{-(1-\alpha)\cdot \ln \widehat{k}(t)} - (n+g+\delta)$$

since $\frac{f(\widehat{k}(t))}{\widehat{k}(t)} = \widehat{k}(t)^{-(1-\alpha)}$ and $x^a = e^{a \ln x}$. Using a first order Taylor Series expansion around the steady state

$$\gamma_{\widehat{k}} \simeq \left(\frac{\partial \gamma_{\widehat{k}}}{\partial \ln \widehat{k}} \mid_{\ln \widehat{k} = \ln \widehat{k}^*}\right)(\ln \widehat{k} - \ln \widehat{k}^*) \tag{2.8}$$

we see that $\gamma_{\widehat{k}}$ is approximately the product of the difference between $\ln \widehat{k}$ and $\ln \widehat{k}^*$ and the derivative of $\gamma_{\widehat{k}}$ with respect to $\ln \widehat{k}$ at $\ln \widehat{k} = \ln \widehat{k}^*$. Differentiating $\gamma_{\widehat{k}}$ with respect to $\ln \widehat{k}$ gives

$$\frac{\partial \gamma_{\widehat{k}}}{\partial \ln \widehat{k}} \mid_{\ln \widehat{k} = \ln \widehat{k}^*} = -(1-\alpha) \cdot s \cdot e^{-(1-\alpha) \cdot \ln \widehat{k}(t)} \tag{2.9}$$

Substituting (2.9) into equation (2.8) and using the steady state condition

$$s \cdot e^{-(1-\alpha) \cdot \ln \widehat{k}(t)} = (n + g + \delta)$$

gives an approximation of $\gamma_{\widehat{k}}$ in the neighbourhood of the steady state

$$\gamma_{\widehat{k}} \simeq -(1-\alpha) \cdot (n + g + \delta) \cdot (\ln \widehat{k} - \ln \widehat{k}^*)$$

Simplifying this to

$$\gamma_{\widehat{k}} \simeq -\lambda (\ln \widehat{k} - \ln \widehat{k}^*) \tag{2.10}$$

where

$$\lambda = (1-\alpha) \cdot (n + g + \delta)$$

we derive an explicit expression of the speed of convergence. It indicates how rapidly the capital stock per effective unit of labour approaches its steady state value.

So far we have concentrated on the steady state value of $\widehat{k}$. In the context of economic growth, however, we are more interested in the steady state of output (per effective unit of labour). However, $\widehat{y}$ approaches $\widehat{y}^*$at the same rate that $\widehat{k}$ approaches $\widehat{k}^*$. This is because, for a Cobb Douglas production function, we can write

$$\ln \widehat{y}(t) = \alpha \ln(\widehat{k}),$$

and

$$\frac{d\ln\widehat{y}(t)}{dt} = \alpha\frac{d\ln\widehat{k}(t)}{dt}$$

and substituting these expressions into (2.10) we verify that the speed of convergence of $\widehat{y}$ is the same as for $\widehat{k}$

$$\gamma_{\widehat{y}} = \frac{d\ln\widehat{y}(t)}{dt} \simeq -\lambda(\ln\widehat{y}(t) - \ln\widehat{y}^*) \tag{2.11}$$

The solution to this differential equation is[4]

$$\ln\widehat{y}(t) = (1 - e^{-\lambda t})\ln\widehat{y}^* + e^{-\lambda t}\ln\widehat{y}(0)$$

Subtracting $\ln\widehat{y}(0)$ from both sides the model suggests a growth equation of the following form

$$\ln\widehat{y}(t) - \ln\widehat{y}(0) = (1 - e^{-\lambda t})\ln(\widehat{y}^*) + (e^{-\lambda t} - 1)\ln\widehat{y}(0) \tag{2.12}$$

Substituting from equation (2.7) for the steady state income per effective unit of labour we derive

$$\begin{aligned}\ln\widehat{y}(t) - \ln\widehat{y}(0) &= -(1 - e^{-\lambda t})\ln\widehat{y}(0) + (1 - e^{-\lambda t})\frac{\alpha}{1-\alpha}\ln(s) \\ &\quad -(1 - e^{-\lambda t})\frac{\alpha}{1-\alpha}\ln(n + g + \delta)\end{aligned} \tag{2.13}$$

Alternatively, we can express this equation in per worker terms instead of per effective unit of labour terms. A per worker formulation will also enable us to test the model empirically. In order to reformulate equation (2.13) recall that we can express output per effective unit of

[4]Please refer to the appendix for further details.

labour as

$$\ln \widehat{y}(t) = \ln\left(\frac{Y(t)}{A(t)L(t)}\right) = \ln\left(\frac{Y(t)}{A(0)e^{gt}L(t)}\right)$$

or

$$\ln \widehat{y}(t) = \ln\left(\frac{Y(t)}{L(t)}\right) - \ln A(0) - gt \tag{2.14}$$

Substituting this expression for $ln\ \widehat{y}(t)$ into equation (2.13) provides us with an expression for growth in output per worker

$$\begin{aligned} \ln y\,(t) - \ln y\,(0) \quad = \quad & -(1-e^{-\lambda t})\ln y(0) + (1-e^{-\lambda t})\ln A(0) + gt \\ & +(1-e^{-\lambda t})\frac{\alpha}{1-\alpha}\ln(s) - (1-e^{-\lambda t})\frac{\alpha}{1-\alpha}\ln(n+g+\delta) \end{aligned} \tag{2.15}$$

where $y(t) = \frac{Y(t)}{L(t)}$ stands for output per worker. In the textbook Solow model the growth rate of output per worker depends on initial output per worker, $y(0)$, the initial level of technology, $A(0)$, the rate of technological progress, g, the savings rate, s, the growth rate of the labour force, n, the depreciation rate, δ, the share of capital in output, α and the speed of convergence, λ. Thus, the model predicts that a high savings rate will affect income growth positively, whereas high labour growth (corrected by the rate of technological progress and the rate of depreciation) will have a negative effect on the growth of income per worker. Furthermore, the model predicts that the coefficients on the savings rate and the growth of the workforce are of the same magnitude but of opposite signs.

Equation (2.15) gives the basic model for growth in output per worker, and has been used as the framework for empirical analysis in MRW, Islam (1995) and CEL.

Using data on growth rates averaged over the period 1960-85 in a cross section regression for 98 countries MRW find that all coefficients are statistically significant, have the correct

signs and the regression offers a good fit. The restriction that the coefficients on $ln(s)$ and $ln(n+g+\delta)$ are equal in magnitude and opposite in sign is not rejected. However, they suggest that the results are not supportive of the Solow model, because the regression implies a value of the share of capital in output of 0.59, which contradicts the conventional assumption of about 0.3. These results are broadly confirmed by Islam (1995). In contrast to MRW he estimates the predictions of the Solow model in a dynamic panel framework. However, CEL reject the textbook Solow model on two accounts. Using dynamic panel data and estimating the textbook Solow model by first differenced GMM estimation, they reject the restriction on the coefficients on $ln(s)$ and $ln(n+g+\delta)$. In addition they find an implausibly *low* value of the share of capital in output of about 0.10.

2.2.2 The Augmented Solow Model

Since MRW find that their results are not supportive of the textbook Solow model, they suggest that augmenting the Solow model by including human capital as a fourth factor of production can improve the predictive power of the model. Thus, the production function becomes [5]

$$Y(t) = K(t)^{\alpha} H(t)^{\beta} (A(t)L(t))^{1-\alpha-\beta} \tag{2.16}$$

or in the intensive form

$$\widehat{y}(t) = \widehat{k}(t)^{\alpha} \widehat{h}(t)^{\beta}$$

where $H(t)$ is the stock of human capital and $\widehat{h}(t) = \frac{H(t)}{A(t)L(t)}$ denotes human capital per effective unit of labour. We assume that there are decreasing returns to physical as well as human

[5] Since the model is very similar to the textbook Solow model, we will keep the discussion relatively brief and concentrate on the results.

capital, thus, $\alpha + \beta < 1$. A constant fraction of output is invested into physical capital, s_k ,and a constant fraction of output is invested into human capital, s_h . Analogous to the previous section the equations of motion for both types of capital are

$$\dot{\widehat{k}}\,(t) = s_k \widehat{y}(t) - \widehat{k}(t)(n + g + \delta) \tag{2.17}$$

and

$$\dot{\widehat{h}}\,(t) = s_h \widehat{y}(t) - \widehat{h}(t)(n + g + \delta) \tag{2.18}$$

These equations of motion are based on the assumption that physical as well as human capital depreciate at the same rate, δ. These two equations allow us to determine the steady state physical and human capital

$$\widehat{k}^* = \left(\frac{s_h^{\beta} s_k^{1-\beta}}{(n + g + \delta)} \right)^{\frac{1}{1-\alpha-\beta}}$$

and

$$\widehat{h}^* = \left(\frac{s_k^{\alpha} s_h^{1-\alpha}}{(n + g + \delta)} \right)^{\frac{1}{1-\alpha-\beta}}$$

The steady state output per effective unit of labour can be determined by taking logarithms of the production function

$$\ln \widehat{y}(t) = \alpha \ln \widehat{k}(t) + \beta \ln \widehat{h}(t)$$

and substituting the steady state values for $\widehat{k}^*$and $\widehat{h}^*$

$$\begin{aligned} \ln \widehat{y}(t) \;\; = \;\; & \frac{\alpha}{1-\alpha-\beta} \ln(s_k) + \frac{\beta}{1-\alpha-\beta} \ln(s_h) \\ & + \frac{\alpha+\beta}{1-\alpha-\beta} \ln(n + g + \delta) \end{aligned}$$

A log linear approximation around the steady state output per effective unit of labour gives as in the textbook Solow model

$$\gamma_{\widehat{y}} = \frac{d \ln \widehat{y}(t)}{dt} \simeq -\psi(\ln \widehat{y}(t) - \ln \widehat{y}^*)$$

However, the speed of convergence in the augmented Solow model is

$$\psi = (1 - \alpha - \beta) \cdot (n + g + \delta)$$

Growth in output per effective unit of labour can thus be expressed as

$$\begin{aligned} \ln \widehat{y}(t) - \ln \widehat{y}(0) &= -(1 - e^{-\psi t}) \ln \widehat{y}(0) + (1 - e^{-\psi t}) \frac{\alpha}{1 - \alpha - \beta} \ln(s_k) \\ &\quad + (1 - e^{-\psi t}) \frac{\beta}{1 - \alpha - \beta} \ln(s_h) \\ &\quad - (1 - e^{-\psi t}) \frac{\alpha + \beta}{1 - \alpha - \beta} \ln(n + g + \delta) \end{aligned}$$

Formulating this growth equation in terms of per worker values

$$\begin{aligned} \ln y(t) - \ln y(0) &= -(1 - e^{-\psi t}) \ln y(0) + (1 - e^{-\psi t}) \ln A(0) + gt \\ &\quad + (1 - e^{-\psi t}) \frac{\alpha}{1 - \alpha - \beta} \ln(s_k) \\ &\quad + (1 - e^{-\psi t}) \frac{\beta}{1 - \alpha - \beta} \ln(s_h) \\ &\quad - (1 - e^{-\psi t}) \frac{\alpha + \beta}{1 - \alpha - \beta} \ln(n + g + \delta) \end{aligned} \tag{2.19}$$

gives the augmented Solow model for growth in output per worker. In addition to the textbook version, the augmented Solow model predicts that growth of output per worker also depends on human capital investment. The sum of the coefficients on investment in physical capital,

labour force growth and investment in human capital should be zero. Equation (2.19) has been used as the framework for empirical analysis in MRW, Islam (1995) and CEL.

MRW find that all coefficients in the augmented Solow model are statistically significant and of the right sign. Furthermore, the share of capital is much closer to the traditionally assumed 0.3. Thus, MRW conclude that adding human capital to the Solow model improves its performance.

However, using panel data Islam (1995) finds that the human capital measure is not significant in the OLS levels regression and significant but negative when he uses a within groups estimator. Thus, he concludes that the empirical test does not provide much support for the augmented Solow model.

Using a first differenced GMM estimator CEL find that they cannot reject the hypothesis that the coefficients on savings, growth of the workforce and their human capital measure sum to zero. However, their coefficient on human capital is also significantly negative. This result leads them to reject the augmented version of the Solow model.

Hence, in the literature we find mixed support for the (augmented) Solow growth model. However, Bond, Hoeffler and Temple (1998) argue that the results obtained by CEL as well as by Islam (1995) are likely to suffer from serious small sample biases. They suggest the use of the Blundell and Bond (1998a) system GMM estimator in a dynamic panel data framework. We explore these estimation issues in more detail in Chapter 3, where we present our empirical tests of the (augmented) Solow model.

2.2.3 Convergence

In the Solow model an economy will converge to its steady state no matter from which initial capital stock and output it starts from. In the empirical growth literature, however, there is a strong emphasis on the examination of convergence in a cross section of economies, rather than on the analysis of how a single economy converges to its steady state. Within the framework of these cross sectional empirical studies two concepts of convergence are discussed. First, the concept of absolute convergence, which means that poor countries will tend to experience higher growth rates in per capita income[6] than rich countries, irrespective of any other characteristics the country may have. Second, the concept of conditional convergence, which says a country experiences higher growth rates the further it is from its steady state. We will now discuss each of the two concepts in turn.

Absolute Convergence

As discussed above, in the Solow growth model each country converges to its steady state level of per capita income and the speed of convergence is inversely related to the distance from the steady state. Consider a group of economies with identical structural parameters for preferences and technology. Thus, according to the Solow model these countries will converge to the same levels of income and asymptotic rates of growth. The countries starting of with lower per capita income are further away from the steady state and will thus converge to the steady state at a higher rate than initially richer countries. The main reason for these different transitional growth rates are the diminishing returns to capital. The convergence to the same steady state

[6]The Solow model provides a framework for the analysis of growth in *output per worker*. However, many empirical studies examine the growth in *income per capita*. Income and output can only be used synonymously if we abstract from net foreign assets. Many empirical studies use income per capita rather than per worker data, probably because per worker data is less readily available and often the data quality is poor.

per capita income and asymptotic rates of growth holds if the economies are closed, but factor mobility re-enforces this mechanism. Theory predicts that factors will move across borders to countries with the highest returns for this factor, therefore equalizing capital-labour ratios and factor prices. Even if the structural parameters are not the same across countries, factor mobility could generate absolute convergence. Thus, the Solow model predicts that initially poorer countries 'catch up' with initially richer countries.[7]

Baumol (1986) investigates whether initially poorer countries experienced higher growth rates than initially richer countries by regressing the long run growth rate on per capita GDP measured in 1870. The coefficient on initial GDP per capita is significantly negative. Thus, the study concludes that initially poorer countries experienced higher growth rates and that poorer countries 'caught up' with richer countries over the period 1870 to 1979. De Long (1986) challenges Baumol's convergence result on two counts. First, the study is plagued by an *ex post* sampling bias. Baumol's sample includes only countries that are rich and have experienced prolonged economic growth.[8] However, the sample should include countries which *ex ante* seemed likely to converge. Including some more countries in the sample De Long shows that there is no evidence for absolute convergence. Second, De Long argues that initial incomes are measured with measurement error. Measuring initial income with error consequently introduces an error to the growth rate measurement. If initial income is understated, the growth rate is overstated and *vice versa*. Regressing the growth rate on initial income this measurement error causes the coefficient on initial income to be biased toward -1, therefore providing false support for the hypothesis of absolute convergence.

[7] For a more detailed discussion see for example Lucas (1988).

[8] Baumol's (1986) study includes the following sixteen countries: Austria, UK, Switzerland, Belgium, Netherlands, Canada, USA, Denmark, Italy, Austria, Germany, Norway, France, Finland, Sweden, Japan.

An interesting question is whether if absolute convergence holds, the dispersion of per capita incomes decreases over time. This alternative concept of convergence is also referred to as σ-convergence.[9] However, Quah (1993) shows that absolute convergence does not necessarily imply a reduction of the dispersion of per capita incomes across countries over time. He argues that for the arbitrary points in time, t_1 and t_2, the cross-section regression of $Y(t_2)$ on a constant and $Y(t_1)$ is

$$P[Y(t_2) \mid 1, Y(t_1)] = E_c Y(t_2) + \beta(Y(t_1) - E_c(Y(t_1)),$$

where $P[y \mid x]$ is the conditional probability of y given x and

$$\beta = \frac{cov_c(Y(t_2), Y(t_1))}{var_c(Y(t_1))}.$$

Subscript c denotes cross section. Let us assume that the variances are the same at t_1 and t_2 for the cross section (i.e. countries' per capita incomes neither decrease nor increase over time)

$$var_c(Y(t_1)) = var_c(Y(t_2)).$$

According to the Cauchy-Schwarz inequality the following holds

$$\mid cov_c(Y(t_2), Y(t_1)) \mid \leq var_c^{\frac{1}{2}}(Y(t_2)) var_c^{\frac{1}{2}}(Y(t_1)).$$

This implies that $\beta \leq 1$, thus

$$P[Y(t_2) - Y(t_1) \mid 1, Y(t_1)] = \mu - (1 - \beta)Y(t_1).$$

This shows that the coefficient on the initial condition in these regressions is always no greater

[9] Barro and Sala-i-Martin (1995) pp 31-32 and pp 382-387.

than zero, even when the cross section distribution remains invariant over time, i.e. if there is neither convergence nor divergence of per capita incomes across countries over time. Thus, it can be formally shown that absolute convergence does not imply σ-convergence; absolute convergence is a necessary, but not sufficient condition for σ-convergence.

A number of studies show that there is no evidence for the absolute convergence hypothesis in large cross-country samples.[10] In addition a number studies have examined the dispersion of incomes per capita across countries over time.[11] For example, De la Fuente (1997) examines income dispersion for a sample of 118 countries in the period 1960 - 1985. He finds that there was an increase in income dispersion during this period. Furthermore, comparing the distribution of incomes at the beginning and the end of the period, we see that in 1960 the income distribution was unimodal and roughly symmetric, whereas the 1985 distribution is trimodal with a peak slightly above the sample average and the other two peaks close to the tales of the distribution. This suggests that some middle income countries have fallen behind whereas some have been able to 'catch up' with richer countries.

However, there is evidence of absolute convergence in per capita incomes amongst the US states, the prefectures of Japan and a number of European countries. These studies also suggest that the income dispersion decreased over time for the US states and the prefectures of Japan, but that the empirical evidence for European countries is less conclusive.[12] Work by Ben-David (1996) also suggests that there is evidence for the existence of absolute convergence amongst trading partners. Examining the incomes of a core country and its ten major trading partners, he also finds that income disparity decreased over time for the members of this 'convergence

[10] See for example Barro (1991).

[11] For a discussion of the world income distribution dynamics, see for example, Quah (1996), Pritchett (1997) and Jones (1997).

[12] See Barro and Sala-i-Martin (1992) and Barro and Sala-i-Martin (1995, chapter 11).

club'.

Conditional Convergence

In the absence of evidence for absolute convergence across large samples of countries researchers turned to the examination of conditional convergence. According to the Solow model economies will converge to their steady state and the speed of convergence is inversely related to the distance from the steady state. To illustrate the point that poorer countries may no experience higher growth rates consider the following example. Assume that there are two countries with different initial per capita incomes and different savings rates such that the poorer country has a lower savings rate than the rich country. The different savings rates will determine different levels of per capita income in the steady state for the two countries. The rich country with the higher savings rate will have a higher level of per capita income in its steady state. It is theoretically possible that the richer country is proportionally further away from its steady state income than the poor country and thus the richer countries' growth rate would be higher than the poorer countries' growth rate. Thus, the concept of conditional convergence suggests that initially poorer countries will experience higher growth rates conditional on the determinants of the steady state.[13]

In his seminal paper Barro (1991) tests this neo-classical growth hypothesis by regressing the growth rate of income per capita over the period 1960-85 on initial GDP per capita, human capital measures, investment, government spending and other policy indicators. Barro finds that the coefficient on initial GDP per capita is negative and significant.[14] He interprets this

[13] See for example the discussion on conditional convergence in Barro and Sala-i-Martin (1995 pp 26-30).

[14] There are of course empirical cross-country studies of growth that precede Barro (1991). For example Kormendi and Meguire (1981) carry out a 'Barro style' regression using IMF data for 47 countries over the period 1950-77. They also find a negative coefficient for initial GDP.

as evidence for conditional convergence, because holding constant the other regressors included in the growth regressions initially poorer countries will experience higher growth rates. This finding is confirmed in numerous other empirical growth studies, for example by MRW who estimate the (augmented) Solow model.[15] Sala-i-Martin (1997a, 1997b) confirms that the partial negative correlation between growth and initial income is extremely robust.[16]

In summary, there is very little evidence for absolute convergence in the literature. Absolute convergence seems to hold for groups of countries or regions which have similar structural parameters and thus similar steady states. However, as Quah (1993) shows absolute convergence does not imply that the dispersion of per capita incomes decreases over time. A different concept of convergence is the notion of conditional convergence. Initially poorer countries experience higher per capita growth rates on average once we control for the differences in the structural parameters which determine the steady state per capita income levels. There is strong empirical evidence for the presence of conditional convergence.

2.3 Determinants of Economic Growth

In this section we present a discussion of the variables which are most commonly included in growth models.

[15] Barro (1991) proved an enormous stimulus for the research of economic growth. The main reason for this is that he made his comprehensive data set available for other researchers. Thus, it made it easier for other researchers to extend Barro's work. A large number of papers are basically a variation of Barro (1991); researchers added variables they assumed to be of particular importance for the explanation of long term economic growth. For example the 1995 Inter-American Development Bank report 'Overcoming Volatility' examines growth by adding measures of volatility to the basic Barro regression. Further examples include Roubini and Sala-i-Martin (1992) who examine the importance of financial development and trade and the papers presented in the World Bank project 'How do National Policies affect long-run growth?' (published in the Journal of Monetary Policy, Vol 32 (3), December 1993). All of these papers find evidence for conditional convergence.

[16] We discuss these tests for robustness in section 2.4.1 of this chapter.

2.3.1 Human Capital

Theory

The textbook Solow growth model does not account for human capital, but there seems to be some evidence that the augmentation by human capital improves the explanatory power of the model. In the Solow model the returns to factor inputs are assumed to be diminishing and thus, the model cannot easily explain long periods of persistently above average growth of income per capita as observed for Japan, the USA, and some European countries in the past 100 years.[17]

A number of growth models have addressed this issue and see human capital as a central determinant of economic growth.[18] For example Nelson and Phelps (1966) suggest that economic growth is determined by the stocks of physical and human capital. Higher stocks of human capital raises a country's ability to innovate and to absorb new technology.[19] Lucas (1988) on the other hand stresses the importance of human capital accumulation through education and learning by doing. In his endogenous growth model human capital accumulation is an alternative to (exogenously given) technological change which can generate sustained growth.

In order to test whether human capital (investment) can account for economic growth, human capital has to be defined and suitable proxies have to be found.

According to Mincer (1993, p 286) human capital can be defined in the following way:

'Individuals have inherited and acquired abilities, but only the latter clearly differ among countries and time periods. Human capital analysis deals with acquired capacities which are developed through formal and informal education at school and at home, and through training,

[17] Maddison (1982).

[18] For a discussion of endogenous growth and education see, for example, Aghion and Howitt (1998, chapter 10).

[19] Aghion and Howitt (1998, chapter 10) discuss possible extensions of the Nelson and Phelps (1966) approach, and present a model with endogenous technical change.

experience and mobility in the labour market.'

Another way of defining human capital is to compare it with physical capital. As Becker (1993, p 16) points out the main distinction between human and physical capital is the fact that

'you cannot separate a person from his or her knowledge, skills, health, or values the way it is possible to move financial and physical assets while the owner stays put.'

Thus, human capital is embodied in people and education, training and health care are seen as investment in human capital. In contrast to Mincer's definition Becker defines health as a part of human capital.

A lot of theoretical and empirical work in the analysis of human capital is concerned with individual rates of return to investment in human capital.[20] Within the growth debate the question of the social rates of return to investment in human capital is more important. There are likely to be positive externalities to individual investment into human capital, in other words the gain to society might be greater that the sum of the gains to the individuals. Education could, for example, generate the following positive externalities: informed and responsible citizenship, lawful behaviour and improved standards of health. However, social returns are difficult to measure.

The results of investment in human capital are complex. Formal education for example will affect health and the population growth rate through a variety of mechanisms. One obvious effect of education and training is the empowerment of a person to acquire skills and this will make this person more productive. Education will also lead to a healthier life style, more educated people have a longer life expectancy. Health care will have a twofold effect. First, the

[20] The following discussion is mainly based on Becker (1993, chapter 3) and Mincer (1993, chapter 10).

better a person's physical condition is the more productive this person will be. Second, health care results in a longer productive live.

Female education means higher returns to market work for women as if they were not receiving any education and working in the household. Hence, the opportunity costs of child rearing are higher for educated women and they will tend to have fewer children. This in turn will lead to an improvement of human capital, because children of smaller families are healthier and better educated than children of larger families. Although it is clear that investment in human capital will not have instantaneous effects on population growth and productivity, it is not clear how long the time lags between educational improvements and the effects on population growth and productivity are.

To sum up, we expect the effects of human capital investment on economic growth to be positive, because this form of investment might exhibit positive externalities, education and training will make labour inputs more productive, life expectancy will rise, and population growth will be lower. On the other hand the relationship between growth and education is interdependent. Growth also affects investment in human capital and population growth in a number of ways. First, economic growth will lead to urbanization and child rearing in urban areas is more expensive than in rural areas. Thus, family size will show a tendency to decrease in growing economies. Second, economic growth will tend to increase wages. Rising wages will attract people who have so far worked in non-market sectors, in particular in the household and subsistence farming. Women will have a stronger incentive to limit family size, since the opportunity costs of child rearing are now higher. Third, the demand for skilled labour rises which will encourage parents to invest in the quality of their children and not in quantity.

Empirical Evidence

In the first part of this section we concentrate on the discussion of education, followed by a discussion of health.

A large number of studies include a measure of education in their growth regressions. For reasons of data availability and comparability education proxies concentrate on a particular aspect, formal schooling. Thus, informal education, on-the-job training and learning by doing are not examined. In the literature levels of schooling (average years of schooling in the working age population) as well as flows of schooling (enrolment rates) have been used to examine the impact of human capital on growth. It is not clear, however, which type of schooling (primary, secondary, tertiary or total number of years of schooling) should have the largest effect on growth.

Barro (1991) regresses the average annual GDP growth rate over the period 1960-85 on initial GDP and a number of other variables. He includes the initial primary as well the initial secondary school enrolment rates in the same regression and finds that they are both positive and significant at the one percent level. This result is robust to different specifications. However, using the enrolment rates from 1950 instead of 1960 they are not significant. This result is attributed to the poor data quality of the 1950 UN enrolment data. Using adult literacy rates as an alternative measure of human capital confirms the previous result of a positive partial correlation between growth and human capital. Barro attributes particular importance to this human capital result. He concludes '*poor countries tend to catch up with rich countries if the poor countries have high human capital per person, ... but not otherwise.*'[21] Why he assigns such importance to human capital as a necessary requirement for conditional convergence is

[21] Barro (1991, p 437).

not obvious. Other variables, for example his measures of political instability (number of revolutions and political assassinations) and a measure of macroeconomic distortions (a price distortion index) seem also to be robustly correlated with growth in his analysis.

MRW also find a significantly positive partial correlation between their human capital measure and growth. They test the augmented Solow model and use the product of enrolment rates for secondary schooling in 1960 and the fraction of the working age population that is of school age (aged 15 to 19) as a proxy for investment in human capital. On the other hand, testing the same empirical model with panel data Islam (1995), CEL and Bond, Hoeffler and Temple (1998) find that secondary school enrolment rates are not significant.

Levine and Renelt (1992) test the robustness of the partial correlation between growth and primary as well as secondary enrolment rates by Leamer's extreme bounds test. They find that the significance of the enrolment rates depends on the conditioning information set, i.e. the other right hand side variables included in the growth regression. Sala-i-Martin (1997a) argues on the other hand that Leamer's extreme bounds test is too restrictive and he therefore analyzes the entire distribution of the estimates of the coefficient of a particular variable.[22] He finds that the primary school enrolment rate is robustly correlated with growth, while the secondary and tertiary enrolment rates are not. In addition his study finds that human capital measured as the total number of schooling, either total, primary, secondary or tertiary, are not robustly correlated with growth.

However, the insignificance of the enrolment rates may be due to data problems. Gemmell (1996) shows that initial enrolment rates are only a good proxy for human capital if certain

[22] Please refer to section 2.4.1 for a more detailed discussion of the extreme bounds analysis as well as Sala-i-Martin's (1997a) results.

conditions hold, such as constant enrolment rates over time and large population growth and/or a large number of time periods. Empirically enrolment rates have increased over time and thus the human capital stock does not only depend on initial enrolment rates, but also on the growth of enrolment rates. Furthermore, the stock of human capital depends on the initial labour force and its growth. If labour force growth is included in the growth regression, but not the growth in the human capital stock, it is likely that the labour force growth conflates the effects of increases in physical labour and the effect of increases in the effectiveness of labour. Thus, the insignificance of initial enrolment rates as a proxy for human capital does not necessarily mean that there are no effects from human capital on growth.[23] Gemmell (1996) uses UN enrolment data to construct a proxy for initial human capital stock. His empirical results provide some support for the hypothesis that the initial human capital stock as well as the growth in human capital are significant in growth regressions.

Pritchett (1996a) on the other hand does not find that education is significant in growth regressions, he even finds a negative coefficient on human capital in some regressions. He uses two proxies for the stock of human capital. One measure is the educational attainment ratios from Barro and Lee (1994) for the population aged 25-65. The second data is the Nehru, Swanson and Dubey (1995) proxy for the average years of schooling in the labour force which was obtained by using the perpetual inventory method to accumulate enrolment rates. Pritchett (1996a) suggests a number of explanations for his results. First, schooling may not result in an increase in human capital. Second, if the demand for educated labour is stagnant, marginal returns to education are falling. Finally, the institutional environment of many countries may encourage educated workers to join state owned enterprises or the civil service, rather than (the

[23] In the appendix we present a more detailed discussion of Gemmell's (1996) argument.

more productive) private sector. Educated workers may then be mainly involved in wasteful activities such as rent seeking, rather than activities that increase a country's output.

Examining male and female schooling Barro and Lee (1994) find that male secondary school attainment is significantly positive in their growth regressions while female secondary school attainment is significantly negative. However, as Stokey (1994) suggests female schooling may be proxying something different than the influence of female human capital on growth. Male and female school attainment as well as enrolment are highly correlated. It is therefore surprising that both variables are statistically significant in growth regressions. Stokey points out that the female schooling variable becomes insignificant once regional dummies for Africa, Latin America and East Asia are added. This suggests that female education acts as a dummy for those regions in which women receive significantly less formal education than men. Stokey analyzes the difference in male and female education for all countries in the sample. Hong Kong, Singapore, Taiwan and Korea are outliers in the sense that the difference between female and male secondary school attainment lies below two standard deviations from the mean. She concludes therefore that female education acts as a dummy for East Asia.

To summarize, the empirical evidence for a partial correlation between schooling measures and economic growth is mixed. In cross section studies the correlation seems to be strongly influenced by the conditioning information set, i.e. by the other right hand side variables included in the growth regression. In panel data studies schooling measures are found not to be significantly correlated with growth.

We now turn to a brief discussion of a further aspect of human capital, health. It has been widely acknowledged in the literature that health can be a cause as well as an effect of

growth.[24] Healthier individuals may work more intensely, more effectively and longer. Thus, high productivity is a direct result of good health. A large number of studies have examined this issue and Strauss and Thomas (1995, 1998) provide a good overview.

On the other hand income is also likely to be a determinant of life expectancy. Most obviously at the microeconomic level, higher income results in better nutrition and thus better resistance against illness. Richer households can also spend more on health care.[25] At the macroeconomic level wealthier countries are able to spend more resources on public health programmes which reduce mortality rates.

According to a detailed World Bank (1984) study the higher a country's income the higher is life expectancy. Higher incomes are associated with lower mortality rates and since life expectancy at birth measures the number of years a newborn baby could be expected to live if his or her mortality pattern exactly paralleled that of all age groups in the current year, higher income is also associated with higher life expectancy. While this World Bank study emphasises the important role of public health programmes, in particular vaccination and antimalarial spraying, there are other important determinants of life expectancy. Improved infrastructure, in particular better sanitation and access to safe water, as well as (female) education are important determinants of life expectancy.

Pritchett and Summers (1996) examine the health and income nexus by regressing infant mortality rates and life expectancy on income. They prefer to use infant mortality rates as a proxy for health for two reasons. First, infant mortality data is more reliable than life expectancy data, because life expectancy data is often derived by estimation, rather than observed

[24] See for example Behrman (1993).
[25] For a detailed microeconomic theoretical framework please refer to Behrman and Deolalikar (1988).

directly from death registrations. The World Bank, for example, only adjusts its life expectancy data for infant mortality rates to receive new estimates. Second, the causes of infant death and adult deaths are very different. While communicable disease are the main cause of child deaths, the main causes of death for adults are heart and lung disorders and accidents, i.e. there is good reason to assume that high income reduces life expectancy. The authors find income to be positively and significantly related to both health measures. However, the results are stronger in the infant mortality regressions. In addition, using instrumental variable analysis the paper presents evidence that the causal relationship runs from income to infant mortality and life expectancy.

A number of growth studies, for example Barro (1997) and Sachs and Warner (1997) include life expectancy as a regressor without discussing the interdependent relationship between income (growth) and health. They find evidence for a strong positive relationship between life expectancy and growth. Sala-i-Martin (1997a) also finds that life expectancy is robustly correlated with growth. We examine the relationship between growth and life expectancy in more detail in Chapter 4.

2.3.2 Investment

Both neo-classical as well as endogenous growth models are based on a production function which includes capital as a factor of production. Thus, in both types of models investment in the capital stock is one of the key determinants of economic growth. Since investment has such a central role in growth theory the majority of empirical studies of economic growth include a measure of investment in the growth regression. The most popular measure is the real investment to GDP ratio, either measured at the beginning of the period or more often

measured as an average over the entire period.

There is very strong empirical support for the hypothesis that there is a positive partial correlation between investment and growth. A large number of studies including MRW, Barro and Lee (1994) and King and Levine (1994) find the coefficient on investment to be significantly positive in cross-country growth equations. Using Leamer's extreme bounds test Levine and Renelt (1992) find this partial correlation to be robust to different conditioning information sets. Sala-i-Martin (1997a) also presents empirical evidence suggesting a robust link between investment and growth.

For panel data the empirical evidence also suggests a strong link between growth and investment. Islam (1995) as well as CEL and Bond, Hoeffler and Temple (1998) find the coefficient on investment to be positive and significant.

Investment itself depends on various policies such as fiscal policy, political instability, openness to trade, and institutional policies. Thus, we expect there to be direct as well as indirect effects of investment on growth. However, as Levine and Renelt (1992) and Sala-i-Martin (1997a) show the direct effect is strong, because even when other (policy) variables are included in the regression investment remains significant in most regressions.

There are a few studies suggesting that investment is not significant. Sachs and Warner (1995, 1997) find that investment is insignificant in regressions which include their measure of openness to trade. The coefficient on their openness measure is positive and significant.[26] They present evidence that poor trade policies reduce investment. This suggests that openness is already helping to account for the difference in investment ratios across countries, thus, causing the coefficient on investment to be insignificant. Barro and Sala-i-Martin (1995, pp 429-30)

[26] Please refer to section 2.3.5 for a more detailed discussion of this openness measure.

as well as Barro (1997, p 33) also find that the average investment ratio is insignificant in a single cross-country growth regression. However, their regressions include life expectancy as a regressor and we find that growth regressions which include life expectancy are likely to suffer from simultaneity bias and also tend to make the coefficient on investment insignificant. We discuss this in more detail in Chapter 4.

One particular point of interest is which kind of investment has the strongest association with growth. Including the investment to GDP ratio as a regressor does not address this question, because this measure neither distinguishes between private versus public nor between equipment versus structures investment. De Long and Summers (1991, 1992, 1993) include a measure of average equipment investment in their cross-section growth regressions and find that the coefficient on equipment investment is positive and significant in different specifications. Sala-i-Martin (1997a, 1997b) finds that using the De Long and Summers data equipment investment is robustly correlated with growth. Thus, the empirical evidence in the growth literature suggests that there is a strong relationship between equipment investment and growth.

The link between growth and public as well as private investment is examined by Easterly and Rebelo (1993). They examine the partial correlation of general public investment as well as educational, health, infrastructure etc. investment with growth and private investment. The study finds that the partial correlation between general government investment and growth as well as between general government investment and private investment is positive and robust to different specifications. Out of the different categories of government investment, transport and communication investment seem to have a strong positive link with growth. However, when private investment is regressed on transport and communication investment as well as other variables, the coefficient on transport and communication investment is insignificant. This

indicates that while infrastructure investment increases the social returns to private investment it may not raise private investment itself. Infrastructure investment is still significant in the growth regression when private investment is included. Easterly and Levine (1997) also find that their infrastructure measure, number of telephones per worker, has a significant and positive coefficient in their cross-country growth regression. Thus, there is some evidence for a positive link between infrastructure measures and growth. However, the link between other types of public investment like agriculture, education, and health, and growth seems to be less robust. In the results reported by Easterly and Rebelo (1993) the coefficients on those types of investment are not significant in a number of regressions.

2.3.3 Fiscal Policy

There are a number of studies which link taxation, public investment, and other aspects of fiscal policy with economic growth. In many studies the size of government is believed to lower savings and growth through the distorting effects from taxation or government expenditure programmes. Barro (1991) measures government size by government consumption minus education and defense expenditure, because he argues that they are more like public investment that public consumption. The coefficient on this government size measure is significant and negative, thus, confirming the hypothesis that government consumption has a negative effect on growth. Levine and Zervos (1994) use Leamer's extreme bounds analysis to test the robustness of this result. They find that the partial correlation between government consumption and economic growth is not robust, because regardless of the conditioning information set, the partial correlation between the two variables is never significant. This may be due to two reasons. First, if countries are choosing the optimal amount of fiscal expenditures and taxes, the ratio of the

government expenditures to GDP should be unrelated to long run growth. The second possible explanation of this result is that government consumption is an aggregate measure of government size and does not capture the distribution of expenditures, the efficiency of government expenditures, or whether the government size is suboptimal.

Government deficits have many effects on macroeconomic variables, one effect is that deficits may reduce growth. A deficit may reduce growth through a number of direct and indirect channels. The initial effect of a budget deficit is a reduction in national savings. Thus, the supply of loans available is reduced, consequently the interest rate rises and private investment decreases. The increase in interest rates also leads to an increased demand for domestic assets and this in turn causes the currency to appreciate. Thus, foreign goods become cheaper for domestic residents, i.e. imports rise and domestic goods become more expensive for foreigners and exports fall. The rise in imports and the fall in exports will then cause a trade balance deficit.

Levine and Zervos (1994) include the budget deficit in their cross-country growth regressions. Using extreme bounds analysis they show that the partial correlation between the government deficit and long run growth is fragile. In their comprehensive study Easterly and Rebelo (1993) on the other hand find that the central government deficit is a fiscal policy variable which is robustly correlated with growth. The authors run three basic regressions and add a number of different fiscal variables. The coefficient on government deficit is significant and negative in all three regressions. There are several possible explanations for this result. First, high deficits may be a proxy for high public debt, which in turn may signal higher taxes and lower public capital in the future. The other possibility is that large deficits are a symptom of macroeconomic instability which is harmful to economic growth.

Easterly and Rebelo (1993) test the hypothesis that higher income taxes lower the net rate of return to private investment, making investment activities less attractive and hence lowering the rate of growth. However, the authors conclude that the link between taxation[27] and growth is fragile. Most of these taxation measures are not significant and often cause initial income to be insignificant. Including more than one taxation measure reinforces this effect. Since the initial income level and taxation measures seem to be highly correlated it is difficult to disentangle the link between taxation and growth in a Barro growth regression.

Devarajan, Swaroop and Zou (1996) examine the relationship between the composition of public expenditure and economic growth for 43 developing countries. The starting point of their analysis is that according to earlier studies total government consumption is negatively correlated with growth. At the same time there is evidence for a positive relationship between certain types of public investment and growth.

Their regression results suggest that capital expenditure is negatively correlated with growth, whereas current expenditure is associated with higher growth. This result stands in contrast to the conventional wisdom that developing countries lack infrastructure investment and that capital expenditures have to be increased. These results were reversed for a sample of 21 'high-income economies'[28] for a similar regression. For this sample capital expenditure is positively associated with growth, while the relationship between growth and current expenditure is negative. One problem with the developing country sample may be that the efficiency of expenditures are not captured by the data. More capital goods expenditure might not necessarily

[27] Easterly and Rebelo (1993) use four different taxation measures (1) the statutory tax rates, (2) revenue from different types of taxes expressed as a fraction of GDP, (3) income weighted marginal income tax rates, and (4) 'marginal' taxes which were obtained by regressing the revenue from each type of tax on its tax base.

[28] All OECD countries except Greece, Portugal and Turkey, which are not classified as 'high income economies' by the World Bank Development Report 1994.

lead to more capital goods. A further problem may be that operations and maintenance expenditures often make a capital good productive, but they are classified as current expenditures. Some capital investment projects might also not increase productivity, because the investment is used for consumption goods for powerful members of society.

The analysis for more detailed public expenditure categories show that transport and communication are negatively correlated with growth. This result is not surprising keeping in mind that most of infrastructure spending is capital spending. The coefficients on other public expenditure shares such as health, schooling, and defense are insignificant. One problem with examining these more detailed categories of public spending is that the sample size is greatly reduced; from 305 to 54 observations. However, if we believe that these results do not depend on sample size, a possible interpretation is that more expenditure in the education and health sector does not necessarily augment the stock of human capital and thus, does not necessarily induce higher growth.

2.3.4 Monetary Policy

Inflation is considered to cause costs to the economy[29] and thus, many economists view price stability to be beneficial for resource allocation and for economic growth. In contrast to the theoretical literature the empirical evidence seems to suggest that there is no strong link between the average rate of inflation and growth. However, this result may be due to the nonlinear effect of inflation on growth.

Kormendi and Meguire (1985) find that the partial correlation between the average inflation rate and growth is significant and negative. Levine and Zervos (1993, 1994) on the other hand

[29] For a comprehensive overview see for example Garfinkel (1989) or Briault (1995).

find that the coefficient on the average inflation is not significantly different from zero. Sala-i-Martin (1997a) confirms this result. Levine and Zervos (1993, 1994) suggest that the effect of inflation on growth may be nonlinear. They formulate the following two hypotheses. First, the 'high $\pi-$hypothesis', which says that only when inflation is very high normal economic relationships break down and hence result in lower growth rates. Second, the 'desensitize hypothesis'. People in high inflation countries will adjust to the high inflation environment and develop mechanisms for coping with high inflation. Economic growth may then be unrelated to high inflation. However, countries with lower inflation rates will not be desensitized to the effects of inflation and thus in low inflation countries an increase in inflation will tend to lower growth rates. Levine and Zervos (1993, 1994) investigate these hypotheses by including the average inflation rate, a dummy for high inflation countries and an interaction term (the product of average inflation and the high inflation country dummy) in the growth regression. This allows for a different slope and intercept coefficient for high inflation countries. The authors try two possible cut off points for high average annual inflation, 80 and 40 percent. Defining 80 percent as high inflation supports the 'desensitize hypothesis', neither the high inflation dummy nor the coefficient on the interaction term is significantly different from zero. However, the coefficient on inflation is significantly negative, indicating that for low inflation countries higher inflation is associated with lower growth. Defining 40 percent as high inflation on the other hand supports the 'high π-hypothesis', high inflation lowers growth in high inflation countries, but not in moderate inflation countries. Barro (1997) also presents some evidence for a non-linear relationship between inflation and growth. He defines low inflation as below 15 percent and finds that the relationship between growth and inflation is insignificant for low inflation countries. For high inflation countries the relationship between inflation and growth

is significantly negative. However, the significance of inflation in growth regressions depends on the definition of 'high inflation' and the *ad hoc* definition by Levine and Zervos (1993, 1994) and Barro (1997) cannot be used to explore the non-linear link between inflation and growth in sufficient detail.

In contrast, Sarel (1996) uses the following approach to derive a definition of 'high inflation'. Using five year period averages the rate of growth is regressed on initial GDP per capita, a number of other explanatory variables, the natural logarithm of inflation and an interaction term of a dummy taking the value one if inflation is 'high' times the difference between inflation and the critical 'high inflation' threshold. This regression is run for different values of 'high inflation' and the rate of inflation for which R^2 is maximized is subsequently defined as the threshold value of 'high inflation'. This threshold is eight percent. The coefficient on inflation is insignificant while the interaction term is negative and significant. The estimated coefficient on inflation for high inflation countries (the sum of the coefficient on average inflation and the interaction term) is also negative and significant at the one percent level. Thus, the study concludes that a significant structural break occurs at an average annual inflation rate of eight percent. Inflation below eight percent does not have an effect on growth, while an inflation rate above has a negative impact on growth. Although this study takes great care to define 'high inflation' the method used may cause the results to be biased. Sarel uses a within groups estimator allowing for country and year effects. Such estimations are likely to suffer from short sample and simultaneity biases as we will discuss in more detail in Chapter 3. Hence, it is possible that the structural break of eight percent is a biased estimate.

Besides the actual rate of inflation, the changes in the rate of inflation may have an effect on economic growth. Inflation rate volatility might be an important source of uncertainty and

instability in an economy. The Inter-American Development Bank (1995) examines various measures of volatility and their impact on growth by adding them to the original Barro (1991) regression. The volatility of monetary policy is measured as the log of the standard error from an AR1 process on the growth rate of nominal M1. The coefficient of this volatility measure is negative, but in three different regression specifications it is only significant in one regression. Thus, there is no robust correlation between this volatility measure and economic growth. Sala-i-Martin (1997a) also finds measuring inflation volatility by its standard deviation that volatility is not significant. However, this result may be due to several reasons. One reason might be that the volatility measure is taken from the entire period 1960 to 1985 (or 1990). An examination of this measure in subperiods would probably show that volatility increased for many countries over time. Thus, using a panel analysis would probably be more appropriate to examine the relationship between monetary policy volatility and growth.

2.3.5 Trade Policy

A large number of empirical studies identify openness to international trade as an important determinant of growth. However, amongst the many theoretical models there is no consensus view on the link between trade and growth.[30] As for example discussed by Grossman and Helpman (1991, chapter 9), trade may have an impact on growth through a variety of channels. Due to increased competition x-inefficiencies will be reduced and this will in turn increase economic growth. Furthermore, international trade creates new channels of communication that facilitate the diffusion of technological progress.[31] Lucas (1988) presents a model with learning by doing externalities in which openness can have a strong positive effect on growth.

[30] For a discussion see for example Aghion and Howitt (1998, chapter 11).
[31] For a discussion of technological change and trade see for example Krugman (1995).

On the other hand Grossman and Helpman (1991) construct models of dynamic comparative advantage in which countries have different stocks of knowledge and endowments. Allowing for international knowledge spillovers, the countries' specializations will be according to their endowments. Grossman and Helpman (1991) show that in this model the effect of international trade on the growth rates of each economy's rate of growth is ambiguous. An example of a model in which trade may have a negative impact on growth in developing countries is the work by Young (1991, 1993). Assuming national knowledge spillovers and bounded learning by doing, developing countries are likely to specialize in the production of goods which have exhausted learning by doing effects. As a result trade is likely to inhibit growth in developing countries.

In contrast to the theoretically ambiguous relationship between trade regime and growth, a large number of empirical studies suggest a positive relationship between openness and growth[32]. However, openness to trade has been measured in many different ways and the trade and growth studies are based on a large number of different measures. One problem is that trade policy is difficult to measure directly. Proxies for trade restrictions can be measured in two ways. First, incidence based measures which aim to proxy for trade barriers directly, such as by counting the occurrence of tariff and non-tariff barriers. Second, outcome based measures, such as the trade to GDP ratio, price distortions and black market exchange rate premiums.

Ratio of Trade to GDP

The simplest outcome based measure is the ratio of exports to GDP. The logic behind the use of this openness measure is that protectionist policies and an anti-export bias tend to reduce the

[32] For an overview see Edwards (1993).

share of exports in GDP. However, this measure is problematic for a number of reasons. Mainly, a country might be very open or closed for reasons not related to the trade policy regime; for example large countries tend to have a lot of interregional trade rather that international trade so that they appear more closed than smaller open economies. Furthermore, the distance to markets and endowments will also determine how much a country will trade. Frankel and Romer (1995) regress the trade ratio (exports plus imports divided by GDP) on the physical size of the country, population and other variables and find that country size and population have a significant negative effect on openness. However, the negative effect of population on openness is much weaker than the effect of the physical size.

The empirical evidence that openness measured as the export to GDP ratio is positively related with growth is very weak. Roubini and Sala-i-Martin (1992) obtain a positive and statistically significant coefficient. In contrast to this result Levine and Renelt (1992) as well as Sala-i-Martin (1997a) find that the average export to GDP ratio is not significant in most regressions. Comparing the trade share with other measures Harrison (1996) concludes that the trade share is one of the least robust measures of openness.

Trade Intensity

Leamer (1988) measures openness by trade intensity ratios. The basic idea is to measure openness by the ratio of exports plus imports divided by GNP. However, this ratio may be determined by factors, such as factor endowments, population and distance from markets that are independent of trade policy. In order to take these other factors into account Leamer predicts a country's trade based on a Heckscher-Ohlin model. The difference between actual trade and predicted trade provides a measure of the impact of trade barriers. Dividing this

difference by GNP acts as an adjustment for country-size. A country is said to be 'open' ('closed') if its trade is greater (smaller) than predicted. One problem with this measure is that the residual in the trade regression may be large due to an omitted variable problem and/or measurement problems. Although Leamer's trade model takes many important trade determinants into account, the model cannot account for unobservables, such as taste and technology. Furthermore, there are inherent limitations in measuring factor endowments.

The construction of this index is relatively involved. Trade is divided into resource, agricultural and manufactured trade and 182 different commodity groups are examined. The trade intensity ratios and other related indices as suggested by Leamer are only available for one year, 1982. Edwards (1998) examines nine different measures of openness and finds that Leamer's index is insignificant in growth regressions.[33]

The difference between actual and estimated trade to GDP ratios is also used by Romer and Frankel (1995) to measure openness. Instead of using a Heckscher-Ohlin model, they base their estimation on a gravity model. In their 'pure geography' model they regress the actual bilateral trade ratio for 1985 on the population and physical size of the country and its trade partners, a dummy for landlocked countries and a dummy for a common border between the two countries. These geographical factors are important determinants of trade ($R^2 = 0.36$). By adding a suitable instrument for the countries' income into the equation the explanatory power of this equation is even higher ($R^2 = 0.71$). The residuals of these equations are interpreted as a measure of openness and are included in various cross-country growth regressions as an explanatory variable. The coefficient on this openness measure is positive and significant.

[33]Wolf (1993) constructs a measure of openness which is similar to Leamer's. Edwards (1998) finds Wolf's index to be insignificant in growth regressions.

However, Romer and Frankel (1995) regress the average growth rate of 1960-85 on this 1985 measure of openness. This seems problematic, because openness in 1985 may be the result of income growth in the preceding 25 year period.

Black Market Exchange Rate Premium

A very popular measure of trade distortions is the black market exchange rate premium. Black foreign exchange markets emerge when exchange controls are set in a way that the official exchange rate is below the equilibrium one. The difference between the official and the black market exchange rate is the black market exchange rate premium. The black market exchange rate premium has a distortional effect, because the black market exchange rate premium works as an implicit tax (subsidy). Exporters can either declare their foreign exchange transactions at the official exchange rate or they can smuggle their goods and exchange foreign currency in the black market. Thus, there is an implicit tax on official exports, which is equal to the black market exchange rate premium. Equally, official imports receive an implicit subsidy.[34] A higher black market premium introduces distortions in the economy and is thus expected to lower the steady state level of output per effective worker and therefore reduce the growth rate for given values of the other variables.

The black market exchange rate premium is determined by a number of factors.[35] First, money and demand preferences lead to a particular portfolio of domestic and foreign currency. Second, since the black market exchange rate premium is a tax instrument a higher budget deficit tends to increase the black market exchange rate premium. Third, the terms of trade have an effect on the black market exchange rate premium. An improvement in the terms of

[34] See Claassen (1991, p 17).
[35] For a more detailed discussion see for example Pinto (1991).

trade, i.e. the price of exports in relation to the price of imports rises, will ease the supply of foreign currency and thus the effect of improved terms of trade will be a lower black market exchange rate premium. A further determinant of the black market exchange rate premium is the risk of dealing in the black market. However, an attractive feature of the black market exchange rate premium as an indicator of trade restrictions is that it not only captures the effect of distortions to trade, but also to capital flows and other markets.[36] The use of the black market exchange rate premium as a measure of openness has also been criticized. The black market exchange rate premium is an endogenous variable, it is for example affected by interest rates and penalties for dealing in the black market. This measure might also simply pick up the thinness of the black market.[37]

Therefore, it is not entirely clear which effects are captured by the black market exchange rate premium. However, Barro and Lee (1993), Harrison (1996), Edwards (1998) and Easterly and Levine (1997) find that the coefficient on the black market exchange rate premium is negative and significant. These results support the hypothesis that trade distortions account for a lower growth rate. Levine and Renelt (1992) as well as Sala-i-Martin (1997a) on the other hand find that this partial correlation is fragile since it is found to be not robust to different specifications.

Price Distortion Index

A further possibility of measuring outward orientation is to measure the extent to which the real exchange rate is distorted away from its free-trade level by the trade regime. Dollar (1992) constructs such a real exchange rate distortion index The norm of this index is the price level

[36] Edwards (1992, p 50).
[37] Discussion of Sachs and Warner (1995, p 106).

that corresponds to a country's particular resource endowments. Over- and undervaluation of the real exchange rate is measured relative to this norm. The price distortion index is determined in three steps. First, using data from PWT 5.5 the index of a country's relative price level is determined by multiplying the ratio of the country's price index to the US price index with the nominal exchange rate. This measure of the relative price level is similar to the usual measure of the real exchange rate, except that here the price indices employed have the same weights in each country.

As a second step the relative price level is corrected for different endowments. Dollar (1992) uses GDP per capita as a proxy for the different endowments, and the relative price level is regressed on GDP per capita. The relationship between the relative price level and GDP per capita is positive. To obtain the index of price distortion, actual and fitted values are compared: countries which have data points above the regression line are interpreted as inward oriented, because their price level is higher than it should be. Africa and Latin America are inward oriented regions, whereas Asia is an outward oriented region. Finally, the index of distortion is obtained by dividing the actual price level by the predicted price level. Dollar (1992) also computes the variability of the real exchange rate as an index of distortion, this variability index is computed by measuring each country's real exchange rate distortion index around its mean during the period 76-85.

Dollar uses these two indices of trade orientation to examine the relationship between openness and growth in a cross-country regression using data from 95 developing countries from 1976 to 1985. Regressing the average per capita GDP growth on the index of the real exchange rate distortion, the real exchange rate variability and the average investment share in GDP shows that there is a significant negative relationship between these trade orientation measures

and growth. This supports the hypothesis that outward oriented LDCs grow faster.

Edwards (1992) includes this price distortion index in his growth regressions, but its coefficient proves to be insignificant. These results raise the suspicion that Dollar's result may be due to his particular choice of the time period, 1976 to 1985. Using Dollar's price distortion index also raises a number of different issues. The index is based on tradeables and non-tradeables, but for the analysis of the link between openness and growth the index should only be based on the prices of tradeables. Harrison (1996) modifies the index by 'purging' it of its non-tradable component. This measure is not significant in any of her growth regressions. According to Pritchett (1996b) Dollar's price distortion measure suffers from a number of shortcomings. For example only very few African countries have been included in the benchmark comparison for the International Comparison Project on which PWT is based. Thus, price level data from most African countries is likely to be poor. Furthermore, the distortion index measures the overvaluation of a currency relative to its purchasing power parity (PPP) level and overvaluation can be caused by a number of factors which are not rooted in trade policy, e.g. by foreign aid availability.

Tariffs and Non-tariff Barriers

Tariffs are a straightforward measure of openness, since tariffs will invariably restrict trade. However, it has often been argued that (in developing countries) tariff protection is not as important as non-tariff barriers. There are various types of non-tariff barriers, such as licenses, quotas and prohibitions. The legal nature of tariff and non-tariff barriers makes them an easily monitorable indicator of openness, for example UNCTAD provides data on import charges and the coverage of non-tariff barriers. Pritchett (1996b) finds that non-tariff barriers are not

correlated to other measure of openness, the correlation coefficients are throughout insignificant and very often display the wrong sign. However, Edwards (1998) includes the average tariff on manufacturing in his growth regression and finds the coefficient to be significant and negative.

Non-tariff barriers on the other hand do not seem to be a good measure of openness. Pritchett concludes that the reason for the bad performance of this openness measure is the generally discretionary nature of non-tariff barrier implementation. Thus, the removal of a particular type of legal restriction does not necessarily increase the outward orientation of a country. Edwards (1998) seems to confirm this hypothesis, non-tariff barriers were not found to be significant.

Multivariate and Bivariate Measures of Openness

The World Bank (1987) categorized a sample of developing countries according to their openness. Either countries are strongly outward oriented, moderately outward oriented, moderately inward oriented, or strongly inward oriented. The classification is made on the basis of various measures of trade policy, tariffs, subsidies and quantitative restrictions. Although considerable analytical effort has been made in deriving this classification, it is nevertheless a subjective measure of the orientation of the trade regime. However, Roubini and Sala-i-Martin (1992) as well as Edwards (1998) find this trade orientation measure to be significant.

Sachs and Warner (1995) construct a binary measure of openness. They classify a country to have a closed trade policy if it has at least one of the following characteristics (1) nontariff barriers covering 40 percent or more of the trade, (2) average tariff rates of 40 percent or more, (3) a black market exchange rate that is depreciated by 20 percent or more relative to the official exchange rate, on average, during the 1970s or 1980s, (4) a socialist economic system, and (5) a state monopoly on major exports. Data is available annually from 1950-1995; the

openness dummy takes a value of zero when the country is closed and a value of one when the country is open. Sachs and Warner (1995, 1997) include the percentage of years in which a country was open in a Barro regression for the period 1970 to 1989. This measure is positive and significant; open economies experienced a higher growth rate than closed economies. This result is also confirmed by Edwards (1998). Sala-i-Martin (1997a, 1997b) claims that the Sachs and Warner openness measure is the openness variable that has the strongest partial correlation with growth.

However, the Sachs and Warner (1995) measure of openness is subject to the following criticism. Although their measure of openness is based on a sophisticated and objective method, the openness proxy itself is a bivariate variable. Countries are either closed or open; the setup does not allow for varying degrees of openness, although the openness dummy is derived from a number of continuous variables. Thus, deriving this measure of openness means that information is lost.

Concluding Comment

There is a large number of studies examining the link between trade and growth. Despite ambiguous predictions from various new growth models the empirical studies find that somehow trade policy has an effect on growth. However, it is difficult to measure trade policy and a large number of proxies have been suggested. They range from outcome based measures such as the trade to GDP ratio, price distortion and black market exchange rate premium to incidence based measures such as tariffs and non-tariff controls. Pritchett (1996b) provides a detailed discussion of the different measures. He finds the different measures to be nearly completely uncorrelated in a cross-country data set and concludes that the different measures proxy for

different aspects of trade policy. Harrison (1996) and Edwards (1998) include a variety of different openness measures in their growth studies. Both studies conclude that openness is positive for growth. However, Harrison's growth study differs from others, because she cannot find any evidence for conditional convergence, a well established empirical result. Thus, her results on openness should be treated with caution.

The aim of the openness measure is to capture the impact of trade policy restrictions on growth. However, trade may be restricted due to structural determinants, such as being a landlocked country, and not only due to policy variables. Inclusion of a landlocked dummy provides a check against the possibility that the openness measure picks up geographic trade determinants. Sachs and Warner (1997) find that their openness measure as well as a landlocked dummy are significant in growth regressions.

2.3.6 Institutions

The neoclassical theory of growth and many endogenous growth models do not consider the effect of institutions. In contrast the institutionalist school stresses that institutions matter for economic growth. North (1989, p 1321) defines institutions in the following way: '*Institutions are rules, enforcement characteristics of rules, and norms of behaviour that structure repeated human interaction.*' Basically North builds his argument for the importance of institutions on the assumption that specialization and division of labour as we find it in industrialized economies can only be achieved when transaction costs are low. Institutions, as for example well-specified and well-enforced property rights, will lower transaction costs and thus promote economic growth. The definition of institutions encompasses formal organizational structures and agendas as well as informal structures such as social norms, taboos and standards of

behaviour. For empirical studies some institutions are obviously easier to analyze than others.[38] Barro (1991) includes two measures of political instability in his study. First, the number of assassinations per million population per year and second, the number of revolutions and coups per year. Both variables have significantly negative coefficients, supporting the common sense view that political instability is detrimental to economic growth. Besides controlling for political stability Barro also includes variables for the economic system, he constructs two dummies for socialist and mixed economies. The classification is based on Gastil (1982). Neither of the dummies is significant.

Kormendi and Meguire (1985) also use the Gastil (1982) index to examine the relationship between institutions and economic growth. The Gastil index is an index of political rights which runs from 1 (most democratic) to 7 (least democratic).[39] Kormendi and Meguire construct a dummy variable which takes the value one for countries which have a Gastil index of one or two and which takes the value zero otherwise. Regressing the average growth rate on various determinants this dummy variable is not significant. However, regressing the investment to GDP ratio on the same variables shows that the coefficient of the dummy is positive and significant. This might suggest the existence of a two-stage process, institutional reliability matters for investment and investment promotes growth.

Barro (1997) finds evidence for a non-linear relationship between the Gastil index of democracy and growth. Countries which have a medium level of civil rights experienced the highest growth rates on average, while countries with either very low or very high levels of civil rights experienced lower growth rates on average. Barro (1997) suggests that democratic institutions

[38] For an overview of studies which examine the link between institutions and growth see Aron (1999).

[39] Gastil (1982 p 4) defines political rights as follows: '*Political rights ... are primarily the rights to participate directly or through freely elected representatives in the determination of the nature of law and its administration in a society.*'

provide checks on governmental power and limit the growth retarding effects generated by kleptocrats. On the other hand, high levels of civil rights, as in Western democracies, are very often associated with re-distributive policies and lobbying activities by special interest groups, both of which are also regarded as having a retarding effect on growth.

Knack and Keefer (1995) use different institutional indicators to examine the relationship of institutions and growth. The authors use data from two private international investment risk services: International Country Risk Guide (ICRG) and Business Environmental Risk Intelligence (BERI). ICRG data provides ratings starting in 1982 for the following variables: expropriation risk, rule of law, repudiation of contracts by government, corruption in government, and quality of bureaucracy. From the BERI data, which provides ratings from 1972 onward, the following variables were used: contract enforceability, infrastructure quality, nationalisation potential and bureaucratic delays. In order to avoid multicollinearity problems, the variables were aggregated to just two indices from initial data, called ICRG and BERI. The higher the index is, the stronger the institutions are. The correlation of these indexes with assassinations, revolutions and coups, and the Gastil index is relatively low. This suggests that the ICRG and BERI indices contain different information to the previously used institutional indicators. Both institutional indicators are subsequently included in a growth regression using average data from 1974 to 1989, and both coefficients are positive and significant. This provides some support for the institutionalist view that institutions, in particular property rights, matter for economic growth. Both indicators are also significant in an investment regression supporting the view that there is a direct effect of institutions on growth and an indirect effect through investment. However, the authors do not include investment and the institutional indicators in the same regression, in other words they do not test whether there is a direct and an indirect

effect and whether one dominates the other. There is also another caveat, Knack and Keefer (1995) measured ICRG for 1982 and BERI for 1972. We have no idea how representative these two years are for the entire period. The BERI measure may also be a result of growth in the preceding period.

2.3.7 Summary

In this section we discussed some of the most popular variables which have been included in growth regressions. There is strong evidence that the coefficient on initial income per capita is significant and negative, which has been interpreted as evidence for β-convergence, i.e. poor countries will tend to 'catch up' with rich ones, *ceteris paribus*. The investment to GDP ratio has been found to be robustly correlated with growth in a large number of studies. Our review of the trade policy and growth relationship showed that the evidence is somewhat mixed, but that certain measures of openness, such as the Sachs and Warner (1995) measure and the black market exchange rate premium, are significant in many growth regressions. With respect to fiscal policy the literature finds that a high government consumption to GDP ratio and fiscal deficits are negative for growth. There is some evidence that the relationship between inflation and growth is non-linear. Only high inflation rates are detrimental for growth, with variation at lower inflation rates having no impact on growth. As our review showed, the evidence for a positive effects from human capital and institutions is relatively weak. It is difficult to find suitable proxies for human capital and institutions and the fragility of the results may be due to conceptual problems and data quality.

We want to stress that this literature survey is only selective and that a number of other determinants of growth have been suggested in the literature. For example regional effects seem

to be quite strong. Amongst others Barro (1991) finds that the coefficients on a dummy for Sub-Saharan African countries and on a dummy for Latin American countries is significant and negative. Thus, on average African and Latin American countries experienced lower growth rates, but the growth model is not able to account for this difference in growth rates. Easterly and Levine (1997, 1998) find that when the average growth rate of a countries neighbours is included as a regressor in a 'Barro style' regression, the coefficient on the Africa dummy becomes insignificant. However, we are sceptical as to how far the inclusion of this neighbourhood effect explains Africa's growth performance better, because a neighbourhood variable is very similar to the concept of a regional dummy.

A number of studies examine the link between cultural variables and growth. Easterly and Levine (1997, 1998) find that high ethnic diversity is negative for growth and Barro (1997) and Sala-i-Martin (1997a, 1997b) find that religious affiliations matter for growth. We examine the link between regional as well as cultural variables and growth in Chapter 5.

2.4 Econometric Issues

We now turn to a discussion of three econometric issues. First, as we have seen there is a vast number of growth studies stressing the importance of different variables. We present a discussion of possible sensitivity analysis which examines the robustness of the link between these proposed variables and growth. Second, many of these right hand side variables may not be exogenous and thus, many growth regressions may suffer from simultaneity bias. Third, and related, there is the issue of the appropriate estimation technique for growth regressions.

2.4.1 Sensitivity Analysis

One method to examine the robustness of partial correlations in the extreme bounds analysis (EBA) suggested by Leamer (1983). To examine the relationship between a particular variable and growth, EBA works in the following way. Growth is regressed on a number of variables which are always included in the regression, I, on the variable of interest, Z, and on a subset of variables chosen from a pool of variables identified by past studies as potentially important, M.

$$gy = \beta_{Ij} I + \beta_{Zj} Z + \beta_{Mj} M_j + \varepsilon$$

This model is then estimated for all possible combinations of the M variables. For each model j β_{Zj} and its standard deviation σ_{Zj} are estimated. The definition of the lower extreme bound is the lowest value of β_{Zj} minus two standard deviations and the upper extreme bound is defined as the highest value of β_{Zj} plus two standard deviations. Only if β_{Zj} remains significant and of the same sign at the extreme bounds is the partial correlation between Z and gy said to be robust.

Levine and Renelt (1992) use EBA to examine the robustness of a number of right hand side variables in a Barro growth regression. They use different empirical models and use PWT as well as World Bank data. Their study shows that only very few variables are robustly correlated with growth. Only initial GDP and the investment share are robust in all of their different models. The Africa and Latin America dummies are also robust in all the regressions presented in the paper, but Levine and Renelt (1992) do not further comment on the robustness of these regional dummies. A large number of economic, fiscal and policy indicators were not found to be robustly correlated. Thus, the authors conclude that there is '*not a reliable,*

independent statistical relationship between a wide variety of macroeconomic indicators and growth.'[40] Another way of putting it, is that '*nothing is robust*' (Sala-i-Martin (1997a)).

However, the question is whether EBA is the appropriate test to establish robustness. Sala-i-Martin (1997a and 1997b) points out that if β_{Zj} is only insignificant or changes its sign at the bound in one single regression, EBA concludes that this variable's link with growth is not robust. Sala-i-Martin (1997a and 1997b) therefore finds that EBA is too restrictive and suggests a different robustness testing procedure. Basically, he runs the same regressions as in the EBA, but examines the entire distribution of the estimated β_{Zj} coefficients instead of examining only the extreme bounds. If for example 95 percent of the density function for the estimates of β_1 lies to the right of zero and only 55 percent of the estimates of β_2 lie to the right of zero, Sala-i-Martin suggests that variable 1 is more likely to be correlated with growth than variable 2. Thus, with this method some level of confidence to the robustness of each variable can be assigned, instead of making an either robust or non robust judgement.

Sala-i-Martin (1997a) examines two different models, one of which does not include investment in the set of variables always included in the regressions, I, and another which does. We want to concentrate on the model which includes the investment share in GDP, because according to the empirical literature investment seems to be one of the most robust variables and should thus be included in growth regressions. Sala-i-Martin (1997a) finds that for 17 out of 59 variables the fraction of the cumulative distribution function which lies to the left (or right) of zero is larger than 0.95. Variables which are according to this criterion robustly correlated with growth include: initial GDP, the investment share, initial life expectancy, the Sachs and Warner openness measure, the Africa dummy, the Latin America dummy, the equipment in-

[40] Levine and Renelt (1992, p 943).

vestment share, fractions of Muslim, Confucian, Protestant and Catholic population and some institutional variables, such as a rule of law measure.

Durlauf and Quah (1998) point out that Sala-i-Martin's results are difficult to interpret. If variables which are orthogonal to the variable of interest, Z, are included in the regression by definition this inclusion would have no effect on β_{Zj}. Therefore, it should be possible to generate a large number of regressions with the same significant coefficients without being able to make any inferences about the robustness of the coefficient. Durlauf and Quah (1998, p 48) suggest that it is not clear whether the Sala-i-Martin method really contributes to the analysis of robustness, it may be more informative about the covariance structure of the control variables. Although this is in principle a correct criticism of the Sala-i-Martin method we feel that it is unlikely that he 'constructed' regressions in which the coefficient on the Z variable remains unchanged due to orthogonality with the other regressors. Sala-i-Martin took care to include as many different regressors as possible and ran about four million regressions. This is probably the most comprehensive study of 'Barro style' regressions so far. However, although Sala-i-Martin's (1997a, 1997b) test of robustness is appealing, because it provides us with a measure of how likely it is that the different variables are robustly correlated with growth, the test is based on 'Barro style' regressions which are likely to suffer from omitted variable bias and from endogeneity problems. Hence, one should be cautious before attributing causal significance to these partial correlations.

2.4.2 Endogeneity

Most empirical studies assume (implicitly) that the right hand side variables are exogenous, i.e. that they are uncorrelated with the error term, and that OLS is therefore an appropriate

estimation method. However, if some of the variables are endogenous, i.e. correlated with the error term, OLS will produce biased and inconsistent estimates of the parameters. Such correlation may arise either if shocks to growth directly affect the explanatory variable also (reverse causation), or if some omitted variable jointly determines both growth and the explanatory variable in question. *Per se* there is no reason to assume that for example investment, human capital and population growth are exogenous in economic growth models.

For the purpose of this discussion we will refer to two concepts of exogeneity.[41] First, we will refer to a variable as predetermined if it is uncorrelated with just the contemporaneous and future errors. Second, we will refer to a variable as strictly exogenous if it is uncorrelated with all of the contemporaneous, future and past errors. We can use the following simple simultaneous equations to illustrate these concepts

$$y_t = \alpha_1 x_t + \beta_{11} y_{t-1} + \beta_{12} x_{t-1} + u_{1t} \tag{2.20}$$

$$x_t = \alpha_2 y_t + \beta_{21} y_{t-1} + \beta_{22} x_{t-1} + u_{2t} \tag{2.21}$$

with u_{1t} and u_{2t} mutually and serially uncorrelated. x_t is predetermined for y_t in the first equation if $\alpha_2 = 0$ and x_t is strictly exogenous for y_t if $\alpha_2 = 0$ and $\beta_{21} = 0$.

However, the empirical growth literature seems to have concentrated on the issue of Granger causality. As we will show, Granger causality is not a test for endogeneity. The standard procedure of the Granger causality test is as follows.[42] Variable y_t is regressed on its lagged values and on lagged values of x_t. If the coefficients of the lagged x variables are significant, x

[41] This discussion draws on Maddala (1992, section 9.10). The distinction collapses in the case of a single cross-section regression, but is important in the context of panel data estimation methods.

[42] See for example Pindyck and Rubinfeld (1998, pp 243-244).

is said to Granger cause y. It the coefficients are not significant, x does not Granger cause y. The same exercise is repeated by regressing x on its lagged values and on lagged values of y. If the result is that x Granger causes y and that y Granger causes x, no inferences can be made about causality.

Jung and Marshall (1985) use Granger causality tests to examine the relationship between GNP growth and export growth. Only for four countries in the 37 developing country sample do the tests show that export growth 'causes' GNP growth. For the majority of the countries it was not possible to establish the direction of causality unequivocally. Harrison (1996) also finds that the causality between openness and growth runs in both directions. Blomström, Lipsey and Zejan (1996) examine the relationship between growth and the investment share and they conclude that investment does not Granger cause growth, but that growth Granger causes investment.

However, Granger causality simply means that one variable precedes another and as Jacobs, Leamer and Ward (1979) show, Granger causality tests are not tests for exogeneity. We can illustrate this point by writing the structural equations (2.20) and (2.21) in the reduced form.

$$y_t = \pi_{11} y_{t-1} + \pi_{12} x_{t-1} + v_{1t}$$

$$x_t = \pi_{21} y_{t-1} + \pi_{22} x_{t-1} + v_{2t}$$

If $\pi_{21} = 0$ we conclude that y does not Granger cause x. It does not necessarily mean that x is exogenous. $\pi_{21}$is defined as

$$\pi_{21} = \frac{\alpha_2 \beta_{11} + \beta_{21}}{1 - \alpha_1 \alpha_2},$$

therefore, $\pi_{21} = 0$ implies $\alpha_2 \beta_{11} + \beta_{21} = 0$. From this it does not follow that $\alpha_2 = 0$, i.e. that x

is predetermined nor that $\alpha_2 = 0$ and $\beta_{21} = 0$, i.e. that x is strictly exogenous. Thus, Granger noncausality is a necessary but not sufficient condition for strict exogeneity.

2.4.3 Estimation Methods

In the empirical growth literature the most commonly used estimation method is cross-section estimation as for example in Barro (1991) and MRW. These 'Barro style' regressions use average values for 1960-85 or 1960-90. This approach assumes the production function parameters and initial levels of technology are the same across countries. In particular it is assumed that conditional on other variables, such as investment and population growth, countries' per capita outputs and growth rates converge to the same steady state values. Islam (1995) relaxed the assumption that the initial state of technology is the same across countries. Using averages for five year sub-periods between 1960-85 he constructs a panel data set which consists of 97 countries with a maximum of five observations each. Islam (1995) re-estimates the MRW model and accounts for unobserved country specific effects by using a within groups estimator. Thus, he allows for different intercepts across countries in his estimation. However, as shown by Nickell (1981) the use of the within groups estimator provides biased and inconsistent estimates in a dynamic panel model with a small number of time series observations (T). Furthermore, Islam's (1995) panel estimations, like the 'Barro style' regressions, are likely to suffer from simultaneity bias.

CEL suggest an alternative estimation method for cross-country growth regressions which uses panel data for a relatively large sample of countries but with relatively few time series observations per country. Using the Arellano and Bond (1991) first differenced generalized method of moments (GMM) estimator they address both the small T bias and endogeneity

biases of the within groups estimator. However, as Bond, Hoeffler and Temple (1998) point out, the first differenced GMM estimator as used by CEL can itself be poorly behaved, because lagged levels of the series used by CEL appear to provide weak instruments for subsequent first differences. Bond, Hoeffler and Temple (1998) suggest the use of the Blundell and Bond (1998a) system GMM estimator which has better small sample properties when estimating dynamic models for persistent series.[43]

These cross-section and short T panel approaches maintain the assumption that the slope coefficients, i.e. the production function parameters and the speed of convergence, are common to all countries. A different estimation approach has been suggested by Lee, Pesaran and Smith (1997).[44] They use annual time series data for 102 countries for 1965-89 and relax the assumption of slope parameter homogeneity by allowing for the steady state growth rate and the speed of convergence to differ across countries. Estimating a stochastic Solow model for each country they find evidence for parameter heterogeneity across countries, in particular the steady state growth rates differ across countries.

However, as Temple (1998) points out there are two main problems with the use of annual time series data. First, some variables such as population and schooling are measured only occasionally. The time series data is obtained by interpolation. As a result the time series may be subject to considerable measurement error. Second, when using annual data the short-run variation in growth rates is likely to be dominated by business cycle effects which are neglected in the specification of long run growth models. This suggests that long lags should be included in the regression analysis. However, the relatively short time dimension (about 30 years of

[43] We discuss both GMM estimators in more detail in Chapter 3.

[44] For a discussion of the different estimation methods see also Lee, Pesaran and Smith (1998) and Islam (1998).

data) does not allow for the inclusion of many lags in individual time series regressions without creating a degrees of freedom problem.

Thus, both approaches have their problems. In this thesis we concentrate on some estimation issues in the short T panel approach, maintaining the restrictive assumption that slope coefficients and the rate of technical progress are common to all countries. This approach is widely used, and we believe that our methods offer an important improvement over existing estimation methods. But it is not our intention to suggest that the short T panel approach necessarily dominates the time series approach.

2.5 Summary and Conclusions

This literature review highlights that although the recent interest in economic growth has generated a vast number of empirical studies a number of issues have not been resolved. For example the estimated speed of convergence ranges from about one to ten percent. The literature is also divided about the robustness of the correlation of certain key variables, for example investment and human capital, with growth. A further issue is the importance of country specific characteristics such as regional, geographical and cultural variables.

Most studies concentrate on finding evidence for the particular importance of one variable. For example Barro (1991) and Barro and Lee (1994) concentrate on human capital, Edwards (1992, 1998) and Harrison (1996) on openness, Sarel (1996) on inflation and Easterly et al. (1993) on shocks. It is unlikely that all these variables are equally important for economic growth and it is thus important to eliminate fragile relationships. Levine and Renelt (1992) use Leamer's extreme bounds analysis and find that virtually none of the variables suggested in the empirical growth literature is robustly correlated with growth. Sala-i-Martin (1997a,

1997b) rejects Leamer's extreme bounds test as too restrictive and examines instead the entire distribution of the parameters of interest obtained from Leamer's approach. However, the regressions used for this testing procedure are cross-section OLS regressions and are therefore likely to suffer from biases due to omitted variables and from endogeneity.

Most studies do not address the issue of endogeneity. Some studies suggest the use of Granger causality tests to establish the direction of causality in growth regressions. However, it is worth noting that Granger causality tests are not synonymous with testing for exogeneity. We therefore suggest the use of panel data and GMM estimation methods in order to address the potential bias problems in cross-country growth regressions.

2.6 Appendix

2.6.1 Solution to Differential Equation (2.11)

The differential equation

$$\frac{d \ln \widehat{y}(t)}{dt} = -\lambda(\ln \widehat{y}(t) - \ln \widehat{y}^*) \tag{2.22}$$

where $\ln \widehat{y}^*$ is a constant, has the general solution

$$\ln \widehat{y}(t) = Ae^{mt} + Be^{-nt} + z. \tag{2.23}$$

Taking the derivative of equation (2.23) and setting it equal to (2.22) we find that

$$\begin{aligned} z &= \ln \widehat{y}^*; \\ m &= -\lambda; \text{ and} \\ n &= \lambda. \end{aligned}$$

We can now reformulate equation (2.23)

$$\ln \widehat{y}(t) = Ae^{-\lambda t} + Be^{-\lambda t} + \ln \widehat{y}^*. \tag{2.24}$$

Evaluation at $t = 0$ and $t = \infty$ gives

$$A + B = \ln \widehat{y}(0) - \ln \widehat{y}^*$$

Using these values for A and B we derive the following solution for the differential equation (2.22)

$$\ln \widehat{y}(t) = (1 - e^{-\lambda t}) \ln \widehat{y}^* + e^{-\lambda t} \ln \widehat{y}(0).$$

2.6.2 Human Capital: Stock versus Flow Measures

A number of different education measures have been used as proxies for human capital in growth models. The two most popular measures are school enrolment rates, which is a flow measure, and average years of schooling, which is a stock measure. The two measures are closely related, because the average years of schooling are the sum of past schooling investments. However, as Gemmell (1996) shows, the approximation of stocks by using flow measures may make an interpretation of the results difficult. The approximation of the human capital stock by the school enrolment rate is only justified when specific assumptions are fulfilled. Gemmell (1996) argues that if labour inputs are endowed with two types of skills, physical skills, L, and human capital skills, H, acquired through schooling the stock of human capital at time t, H_t, can be characterized by

$$H_t = H_{t-1} + \alpha_t N_t - \beta_t R_t \tag{2.25}$$

Where N_t stands for labour force, α_t, is the proportion of new labour force entrants, and β_t is the proportion of the retirements, R_t. Expressing the net additions to the labour force as

$$dL_t = N_t - R_t$$

and substituting for N_t in equation (2.25) gives

$$H_t = H_{t-1} + \alpha_t dL_t + (\alpha_t - \beta_t)R_t$$

Approximating α_t by the school enrolment rate and assuming that the human capital stock at time T is the initial stock of human capital plus the accumulated investments in schooling, H_T can be expressed as

$$H_T = H_0 + \sum_{t=1}^{T} \alpha_t dL_t + \sum_{t=1}^{T} (\alpha_t - \beta_t)R_t$$

Since β_t is unobservable[45] H_t is approximated as

$$H_T \approx H_0 + \sum_{t=1}^{T} \alpha_t dL_t$$

Assuming $H_0 = 0$

$$H_T \approx \sum_{t=1}^{T} \alpha_t dL_t$$

and if $\alpha_t = \alpha_0$ for all $t = 1, ..., T$ then

$$H_T \approx \alpha_0 \sum_{t=1}^{T} dL_t$$

[45] Assuming a working life of forty years the following approximation could be made $\beta_t = \alpha_{t-40}$. Gemmell (1996, p 13).

using $\sum_{t=1}^{T} dL_t = L_T - L_0$

$$H_T \approx \alpha_0(L_T - L_0)$$

and if the labour force grows exponentially at rate n,

$$L_t = L_0 e^{nt}$$

H_T can be written as

$$H_T = \alpha_0(L_0 e^{nt} - L_0)$$

and the human capital stock per capita, h_T, is

$$\frac{H_T}{L_T} = h_T \approx \alpha_0 \frac{(L_T - L_0)}{L_T} = \alpha_0 \left(1 - \frac{L_0}{L_T}\right) = \alpha_0 \left(1 - \frac{L_0}{L_0 e^{nT}}\right) = \alpha_0(1 - e^{-nT}) \ .$$

This expression shows that the human capital stock per capita can only be approximated if school enrolment rates are constant over time and if n and/or T are large.

However, enrolment rates have displayed a systematic tendency to increase over time. Thus, the approximation of the human capital stock by using enrolment rates does not seem to be justified. Furthermore, the use of this approximation is critical if a growth regression includes labour growth and school enrolment rates separately. The labour growth variable is then likely to capture the effects of both increases in physical labour and some of the increase in human capital, because for given school enrolment rates, cross-country differences in labour growth imply differences in human capital stocks across countries. Thus, the insignificance of the coefficient on human capital variables therefore might not imply that there are no effects from human capital on growth.

Chapter 3

Panel Data Estimation and the Solow Growth Model

3.1 Introduction

In spite of the large number of endogenous growth models developed over the last ten years, the neo-classical growth model based on Solow (1956) remains a popular model in the debate about the empirics of economic growth. As discussed in the previous chapter, Solow's theoretical framework provides the basis for a number of recent empirical studies, such as Mankiw, Romer and Weil (1992) (henceforth MRW), Islam (1995) and Caselli, Esquivel and Lefort (1996) (henceforth CEL). This chapter has three main parts, a discussion of the theoretical issues, the econometric methodology and the empirical results. In the theory section we provide a brief exposition of the textbook Solow model and its augmented version. In the methodology section we discuss different estimation methods for growth regressions. We show that for a dynamic panel data analysis OLS levels as well as within groups estimation are likely to suffer from biases due to endogeneity and unobserved country specific effects. We suggest a recently developed system GMM estimator as our preferred estimation method for growth regressions, which we

find confirmed by the results as presented in the third section. The last section concludes.

3.2 Theoretical Framework

3.2.1 The Solow Model

As discussed in detail in Chapter 2, in the Solow model growth in output per worker depends on initial output per worker, $y(0)$, the initial level of technology, $A(0)$, the rate of technological progress, g, the savings rate, s, the growth rate of the workforce, n, the depreciation rate, δ, the share of capital in output, α, and the rate of convergence, λ. Thus, the model predicts that a high saving rate will affect growth in output per worker positively, whereas high labour growth (corrected by the rate of technological progress and the rate of depreciation) will have a negative effect on growth in output per worker. The basic model is

$$\begin{aligned} \ln y\,(t) - \ln y\,(0) &= -(1-e^{-\lambda t})\ln y(0) + (1-e^{-\lambda t})\ln A(0) + gt \\ &\quad +(1-e^{-\lambda t})\frac{\alpha}{1-\alpha}\ln(s) - (1-e^{-\lambda t})\frac{\alpha}{1-\alpha}\ln(n+g+\delta). \end{aligned} \tag{3.1}$$

In the augmented version of the Solow model investment in human capital is an additional determinant of growth in output per worker

$$\begin{aligned} \ln y(t) - \ln y(0) &= -(1-e^{-\psi t})\ln y(0) + (1-e^{-\psi t})\ln A(0) + gt \\ &\quad +(1-e^{-\psi t})\frac{\alpha}{1-\alpha-\beta}\ln(s_k) \\ &\quad +(1-e^{-\psi t})\frac{\beta}{1-\alpha-\beta}\ln(s_h) \\ &\quad -(1-e^{-\psi t})\frac{\alpha+\beta}{1-\alpha-\beta}\ln(n+g+\delta) \end{aligned} \tag{3.2}$$

where s_k and s_h denote the proportion of output invested in physical and human capital, respectively.

We refer to equation (3.1) as the textbook Solow model and to equation (3.2) as the augmented Solow model. Equations (3.1) and (3.2) have for example been used as the framework for empirical analysis by MRW, Islam (1995) and CEL. We will follow these studies and base our empirical analysis on the same equations.

3.2.2 The Solow Model and Some Growth Theory Issues

Convergence

Our discussion of the (augmented) Solow model in the previous chapter focused on the transitional growth dynamics of one country to its steady state income[1] path. The speed of convergence to the steady state can be estimated in the following way. Recall that the coefficient on the initial income per worker is $-(1-e^{-\lambda t})$ in the textbook Solow model. Setting

$$\gamma = -(1-e^{-\lambda t})$$

the speed of convergence is given by

$$\lambda = -\frac{\ln(\gamma+1)}{t}$$

where t is the length of the time period. Equivalently the speed of convergence for the augmented Solow model can be estimated as

$$\psi = -\frac{\ln(\gamma+1)}{t}$$

[1] As discussed in Chapter 2 the predictions of the Solow growth model are tested using income data and not output data.

Using cross-country data for the estimation of (3.1) or (3.2), a negative coefficient on initial income is interpreted as evidence for β-convergence (conditional convergence), because holding constant the other regressors included in the growth regressions, initially poorer countries will tend to experience higher growth rates. Barro (1991) as well as MRW and numerous other empirical growth studies find evidence for conditional convergence. Sala-i-Martin (1997a, 1997b) finds that this partial negative correlation between growth and initial output is extremely robust.

Studies such as Barro (1991) and MRW are based on the assumption that the initial state of technology, $A(0)$, is the same across countries and thus the estimated speed of convergence gives the average speed at which a country converges to the common conditional steady state income path.

Typically cross-country regressions span the period 1960-1985 or 1960-90 and the speed of convergence is estimated to be between 0.6 and 2 percent *per annum*.[2] This implies a very low speed of convergence, i.e. it takes between 35 and 115 years to close half the gap between initial income and steady state income. Since the speed of convergence is low the average country is not very close to the steady state. However, the regression equations (3.1) and (3.2) are based on an approximation around the steady state. Only if countries are in the neighbourhood of their steady state do these equations provide a useful basis for our empirical research. If on the other hand countries are not in the neighbourhood of their steady state we should instead concentrate on the transition to the steady state.

Allowing for differences in the initial state of technology (for example Islam (1995) and CEL) means allowing for different steady state income paths across countries. The estimated speed

[2]See for example MRW p 426 and p 429.

of convergence gives the speed at which a country converges to its steady state income path.[3] Panel data studies by Islam (1995) as well as CEL have questioned the conventional finding of a low speed of convergence. They basically argue that the typical OLS results obtained from cross-country growth regressions cause the estimated speed of convergence to be biased downwards. According to their panel data estimations the speed of convergence is found to be between 8 and 13 percent. Thus, countries are converging at a much faster speed to their steady states. These results provide better justification for the assumption that countries are close to their steady state income paths. In the next section we will show, however, that Islam's (1995) within groups estimation as well as the first differenced GMM estimation used by CEL are themselves likely to cause the estimated speed of convergence to be biased upward.

Physical and Human Capital

One main assumption of the Solow model is that the savings rate is exogenous and constant over time. Approximation of the savings rate by the investment rate shows that there is variation over time in the investment data for a number of countries. Taiwan, Korea and Singapore for example increased their investment rates during the period 1960-90. A number of African countries on the other hand experienced a decline in investment rates over the same time period.

Investment in human capital, proxied by enrollment rates or years of schooling, also show changes over time in some countries. In a global sample, African countries experienced the strongest growth in these schooling variables.

Thus, investment rates are not constant over time. In our panel data analysis, however, we will allow the investment rate, the population growth rate and the schooling measure to vary

[3]See Durlauf and Quah (1998) for a more detailed discussion of the different concepts of convergence.

over time.

Technology

The production function on which the Solow model is based includes the input $A(t)$, which is often referred to as knowledge, or as an efficiency parameter, i.e. as an indicator of the state of technology. This parameter $A(t)$ may itself depend on a variety of other variables, such as research and development activities, scope for learning by doing, institutions, climate, and resource endowments. We are therefore likely to describe a variety of unobservables by including the parameter $A(t)$, which will inevitably lead to difficulties in the interpretation of any empirical results. In the following discussion we will refer to the parameter $A(t)$ as the level of technology at time t and to its rate of change as technological progress, g.

Commonly the exogeneity of technology and technological progress in the Solow model has been seen as its main shortcoming. Since exogenous technological progress is a core assumption on which the model rests, this criticism lead to an entirely new branch of the theoretical growth literature. Romer's (1986) seminal paper started the debate on growth models which aim to endogenize technological progress.

A further shortcoming of the model is that although technology is a factor of production, it is not remunerated. Given a linearly homogeneous production function with constant returns to scale, Euler's Theorem says that under the assumption of perfect competition(the price of factor inputs equals their marginal products) the sum of total payments to capital and labour is equal to total output.[4] Technological progress is therefore free. However, this problematic assumption cannot be addressed within the Solow model. In the endogenous growth literature models of

[4]For a proof of Euler's Theorem, see for example Chiang (1984), pp 413-16.

Schumpeterian growth, such as Romer (1990) and Aghion and Howitt (1992), endogenize growth by introducing a research and development sector which produces new technology as a result of deliberate and remunerated research activities.

As we will explain in more detail below, empirical tests of the Solow model are based on the assumption that although the initial state of technology $A(0)$ can be different across countries, the constant rate of technological change is common across countries. Thus, all countries enjoy the same rate of technological progress irrespective of their initial state of technology. Closing the initial technology gap through accelerated technological progress is ruled out. The persistence of a technology gap seems to have at least one well cited empirical counter example, the so called 'East Asian miracle'. The World Bank (1993) report for example provides evidence that relatively rapid productivity growth in these economies has led to high and prolonged output growth. However, there is a contrarian view stating that the East Asian growth experience is not 'miraculous'. An increase in labour force participation rates led to rapid growth of the labour force in the four Asian tiger economies, Taiwan, Singapore, South Korea and Hong Kong. At the same time these countries experienced high investment rates. Thus, Young (1994, 1995) argues that the two main determinants of growth in the Solow model, savings and labour force growth, can explain most of the East Asian growth experience. This contrarian view seems to provide some support for the hypothesis that 'catch up' processes are possible even if the rate of technological progress is the same across countries.

Cross-country regressions implicitly assume that both the level of initial technology and the rate of technological progress are common to all countries. Panel data studies, such as Islam (1995) and CEL, implicitly assume that the rate of technological progress is common to all countries, but allow for unobserved differences in the initial level of technology. Whilst this is a

weaker restriction than that required for cross-country regressions to yield consistent parameter estimates, we are aware that it may still be unduly restrictive. Studies of the diffusion of new technology suggest that diffusion is likely to be costly, and take a considerable period of time.[5] This is indeed an important motivation for allowing initial levels of technology to differ across countries. However, if the diffusion of new technology is not costless and instantaneous, we may want to go further and allow for different rates of technological progress in different countries. Nevertheless, in the absence of reliable data on rates of technological progress, we maintain the standard assumption of a common rate of technical change in this study.[6]

Finally the assumption of labour augmenting technological progress deserves some discussion. Barro and Sala-i-Martin (1995) argue that technological progress needs to be labour augmenting for there to be a steady state.[7] The assumption of technological progress to be only labour, but not capital augmenting seems rather restrictive. However, as Barro and Sala-i-Martin (1995, p 55) suggest, for a model which includes Hicks neutral technological progress, i.e. the ratio of marginal products remains unchanged for a given capital to labour ratio in the presence of technological progress, steady state properties similar to the steady state of the Solow model can be derived. In order to have a steady state with Hicks neutral technology the latest vintage of capital must be of higher quality than the previous vintages. Since the steady state properties of the two models are similar we continue by basing our analysis on a production function with labour augmenting technological progress for expositional ease.

[5] For example, Karshenas and Stoneman (1995) find that there are often several decades between the first application of a new technology, and it achieving a 90 percent market penetration.

[6] Freeman(1994) discusses some of the difficulties in measuring technological change.

[7] For the proof see for example Barro and Sala-i-Martin (1995, pp 54-55).

3.3 Methodology and Data

Typically the empirics of long run economic growth have been analyzed in a cross section framework using average data for 25 or 30 years. MRW follow this approach, but other examples are the studies by Kormendi and Meguire (1985), Barro (1991), and Levine and Renelt (1992). Their analysis is based on a regression of the following form

$$g_i = a_0 + b_0 y_i + c_0 x_i + u_i \tag{3.3}$$

where g_i denotes the growth rate of real per capita GDP over a 25 or 30 year period, y_i is the initial level of real per capital GDP, u_i represents an error term and $i = 1, ..., N$ denotes a country index. For empirical tests of the Solow model the regression includes two additional regressors, the investment rate and population growth. However, most empirical studies are based on more general models and include a range of other socio-economic variables. These variables are either initial values or average values over the time period. For the discussion of the methodology it does not make any difference whether we think of the (augmented) Solow model or more general models and for expositional purposes we will therefore concentrate on one regressor, x_i.

We will not use single cross-country regressions for the analysis of the Solow model for a number of reasons. First, reducing time series to a single (average) observation means that not all available information is used. Second, it is very likely that single cross-section regressions suffer from omitted variable bias. Third, one or more of the regressors may be endogenous. Since single cross-section growth regressions potentially suffer from these problems we will instead use a dynamic panel data approach. Within this framework we concentrate on two estimators,

the Arellano and Bond (1991) first differenced generalized method of moments (DIF-GMM) estimator and the Blundell and Bond (1998a) system generalized method of moments (SYS-GMM) estimator. Both GMM estimators address the bias problems encountered in single cross-section regressions, because in a dynamic panel data model we will be able to account for unobserved country specific effects and allow for the endogeneity of one or more of the regressors.

3.3.1 Dynamic Panel Data Analysis

Single cross-section growth regressions are likely to suffer from omitted variable bias. Obvious candidates for these omitted variables are proxies for technology, institutions and resource endowments. Thus, the error term, u_i , in the single cross-country regression (3.3) is likely to contain unobserved variables which influence growth. These unobserved variables are likely to be correlated with some of the observed regressors and in particular any permanent unobserved influences will necessarily be correlated with the initial GDP level. Thus, the OLS estimation of (3.3) is very likely to suffer from an omitted variable bias. However, accounting for unobserved country specific effects is not possible in the single cross-section model.

An alternative approach is to use time series data for each country and take ten or five year averages instead of averages over the entire period of 25 or 30 years. This provides a panel data set for the study of economic growth. In a panel data model we can then explicitly account for permanent unobserved country specific effects, η_i. This provides a panel data model of the form

$$g_{i,t} = a_0 + b_0 y_{i,t-1} + c_0 x_{it} + \eta_i + v_{it} \tag{3.4}$$

where t denotes points in time $t = 2, ..., T$. For five year averages the first time we observe

the data is in 1960, the next observation is in 1965, which corresponds to $t = 2$, the third observation is in 1970, which stands for $t = 3$ etc. ... . g_{it} denotes growth over a five year period, which starts in $t-1$ and ends in t. The first time period for which the growth data is observed is 1960-65, the second period 1965-70, ... and the last period 1985-90. $y_{i,t-1}$ stands for the logarithm of per capita GDP at the beginning of each time period, i.e. in 1960, 1965, ..., 1985. x_{it} is either a period average or the initial value taken at the beginning of each period.

Since the relevant five year growth rate in (3.1) is the logarithmic difference in GDP per worker, we have the following dynamic panel data model

$$y_{it} - y_{i,t-1} = a_0 + b_0 y_{i,t-1} + c_0 x_{it} + \eta_i + v_{it} \tag{3.5}$$

or equivalently

$$y_{it} = a_0 + b_0^* y_{i,t-1} + c_0 x_{it} + \eta_i + v_{it} \tag{3.6}$$

where $b_0^* = (b_0 + 1)$. It is important to note that the typical panel in the study of economic growth then has a small number of time series observations, i.e. T is small. Asymptotic results discussed below focus on the case where N becomes large for fixed T.

As Hsiao (1986) shows, omitting unobserved time invariant country effects in a dynamic panel data model will cause OLS levels estimates to be biased and inconsistent. The lagged dependent variable, $y_{i,t-1}$, is positively correlated with the permanent effects, η_i. As a result the OLS levels estimate of the coefficient $\widehat{b_0^*}$ in the typical growth regression is likely to be biased upward.[8]

An alternative estimation technique which takes account of the unobserved country specific

[8]See Hsiao (1986, pp 76-78). The sign of this bias can be derived unambigously in cases where there are no x_{it} variables, or where all the variables x_{it} are strictly exogenous with respect to $(\eta_i + v_{it})$.

effects is the within groups estimator. For this method model (3.6) is transformed by subtracting out the time series means of each variable for each country. In this transformation process the country specific effects are eliminated, because the country specific effects η_i are invariant over time. The transformed model is then estimated by OLS.[9] However, as shown by Nickell (1981), using the within groups estimator will also provide biased and inconsistent estimates in a dynamic panel model with fixed T. In contrast to the OLS levels estimation, the within groups estimate of the coefficient b_0^* is likely to be biased downward. Consequently, the estimate $\widehat{b_0^*}$ obtained from OLS levels can be loosely interpreted as an approximate upper bound on this coefficient, and the estimate obtained from within groups estimation can be loosely interpreted as an approximate lower bound.

Furthermore, one or more regressors in (3.6) may be correlated with η_i and/or v_{it}. Both problems can be addressed by using a first differenced GMM estimator (Hansen (1982)). In order to obtain a consistent estimate of b_0^* as $N \longrightarrow \infty$ for fixed T we take first differences of (3.6) which eliminates the country specific effects η_i[10]

$$(y_{it} - y_{i,t-1}) = b_0^*(y_{i,t-1} - y_{i,t-2}) + c_0(x_{it} - x_{i,t-1}) + (v_{it} - v_{i,t-1}) \tag{3.7}$$

Since the differenced lagged dependent variable and the differenced error term are correlated OLS estimation of (3.7) will not produce a consistent estimate of b_0^*, even if the regressor, $x_{i,t}$, is strictly exogenous. Thus, valid instruments have to be found for $\Delta y_{i,t-1} = (y_{i,t-1} - y_{i,t-2})$. Assuming that the errors are independent across countries and serially uncorrelated

$$E(v_{it}v_{is}) = 0 \text{ for } s \neq t$$

[9] For further details see for example Hsiao (1986) and refer to the appendix.
[10] See Anderson and Hsiao (1982).

and that

$$E(y_{i1}v_{it}) = 0 \text{ for } t \geq 2$$

then values of y_{it} lagged two periods or more are valid instruments in the first differenced growth equation, since y_{it-2} and earlier values are generally correlated with $\Delta y_{i,t-1}$, but not with Δv_{it}. According to these assumptions $y_{i,t-1}$ is predetermined with respect to v_{it}, i.e. shocks to GDP in one time period are not correlated with initial GDP of this time period, but shocks to GDP in previous periods may be correlated with initial GDP of this time period. For dynamic panel models of economic growth this seems a reasonable assumption since initial GDP is measured at the beginning of each period.

If the other regressor is strictly exogenous

$$E(x_{it}v_{is}) = 0 \text{ for all } s, t$$

then all the past, present and future values of x_{it} are valid instruments in each of the differenced equations, even if the x_{it} are correlated with η_i. However, there are a number of reasons to believe that some of the regressors in empirical growth models, e.g. investment, may not be strictly exogenous. There may be a feedback mechanism where past shocks to GDP are correlated with current investment. Maintaining the assumption that current shocks to GDP are uncorrelated with current investment, this means that

$$E(x_{it}v_{is}) \neq 0 \text{ for } s < t$$

and

$$E(x_{it}v_{is}) = 0 \text{ for } s \geq t$$

Following Arellano and Bond (1991) we can then use values of x_{it} lagged one period or more as valid instruments in the first differenced growth equation, since the x_{it} are predetermined.

It is also straightforward to treat for example investment as an endogenous variable. This means that we are allowing for correlation between current investment and current shocks to GDP, as well as feedback from past shocks to GDP, i.e.

$$E(x_{it}v_{is}) \neq 0 \text{ for } s \leq t$$

and

$$E(x_{it}v_{is}) = 0 \text{ for } s > t \text{ only}$$

In this case, valid instruments in the differenced equations are values of the endogenous x_{it} lagged two periods or more.

However, as Blundell and Bond (1998a) show, estimators relying on lagged levels as instruments for current differences are likely to perform poorly when the series are close to a random walk. Instead they suggest to estimate a system combining two sets of equations. One set of equations are the differenced equations (3.7)

$$(y_{it} - y_{i,t-1}) = b_0^*(y_{i,t-1} - y_{i,t-2}) + c_0(x_{it} - x_{i,t-1}) + (v_{it} - v_{i,t-1})$$

for which we use lagged levels of y_{it} and x_{it} as instruments, as discussed for the differenced GMM estimation. The other set of equations in the system are the levels equations (3.6)

$$y_{it} = a_0 + b_0^* y_{i,t-1} + c_0 x_{it} + \eta_i + v_{it}$$

Introducing the relatively mild initial condition restriction

$$E(\Delta y_{i2}\eta_i) = 0 \tag{3.8}$$

and the restriction

$$E(\Delta x_{it}\eta_i) = 0 \tag{3.9}$$

allows the use of $\Delta y_{i,t-1}$ and Δx_{it} as instruments in the levels equations.[11] Estimators of this type were proposed by Arellano and Bover (1995). Assumption (3.9) allows the level of x_{it} to be correlated with the unobserved country specific effects, η_i, but requires the changes in x_{it} to be uncorrelated with η_i. This is clearly weaker than requiring the levels of x_{it} to be uncorrelated with η_i. Given (3.9), assumption (3.8) will be satisfied provided the process (3.6) has been generating the y_{it} series for a sufficiently long time - this is a kind of stationarity restriction. Alternatively, Blundell and Bond (1998b) show that sufficient conditions for (3.8) and (3.9) are for the series x_{it} and y_{it} to have time-invariant means.

As an empirical matter, the validity of these additional instruments can be tested using standard Sargan tests of overidentifying restrictions, or using Difference Sargan tests comparing first differenced GMM and system GMM results (*cf.* Arellano and Bond (1991)). Exploiting the additional restrictions (3.8) and (3.9) means that the system GMM estimator is more efficient than the differenced GMM estimator, provided these restrictions are valid. Blundell and Bond (1998a) show that the efficiency gain can be dramatic when the series are close to being random walks, and that the differenced GMM estimator can also have large finite sample biases in these

[11] Δx_{it} can be used if x_{it} is strictly exogenous or predetermined with respect to v_{it}. This should be replaced by $\Delta x_{i,t-1}$ if x_{it} is endogenous. The use of further lags beyond Δy_{it} and Δx_{it} can be shown to be redundant, given the moment conditions $E(y_{i1}v_{it}) = 0$ for $t \geq 2$ and $E(x_{it}v_{is}) = 0$ for all t, s.

cases. In a model with no x_{it} variables, the differenced GMM estimate of the coefficient b_0^* is found to be biased downward in small samples when the instruments are weak.

We now turn to a brief discussion of the unobserved country specific effects, η_i. One important question is whether we have to account for these unobserved country specific effects.

If the unobserved country specific effects are not significant, equation (3.6) becomes

$$y_{i,t} = a_0 + b_0^* y_{i,t-1} + c_0 x_{it} + v_{it} \tag{3.10}$$

and we could estimate (3.10) by instrumenting the endogenous regressors in the following way. Assuming as before that the errors are serially uncorrelated, that $y_{i,t-1}$ are predetermined and that x_{it} are endogenous, values of x_{it} lagged one period or more are valid instruments for x_{it} in the levels equations. (Since $y_{i,t-1}$ are predetermined there is no need to instrument for them in the levels equations.) We refer to this estimation method as IV levels.

In the next section we compare the results estimating (3.10) by IV levels and estimating (3.6) by system GMM. (3.10) is a restricted version of (3.6), because according to model (3.10) all countries have the same intercept term, whereas (3.6) allows the intercept to vary across countries. Thus, the comparison of the IV levels and the system GMM results will allow us to examine whether significant country specific effects are present.

3.3.2 Data and Sample

One important question when testing the Solow model is whether to use per capita or per worker variables. According to the Solow model it seems more appropriate to use per worker GDP and the growth of the workforce, because the model is based on a production function and not every person contributes to production. We would expect the results of growth regressions

to be sensitive to the very different dependency ratios across countries. In Sub-Saharan African countries for example the population growth rate has been higher than the growth of the workforce on average.

MRW use per worker variables, whereas Islam (1995) and CEL use per capita values. We have run all regressions in per worker as well as per capita terms. The results do not seem to be sensitive to the choice of per worker or per capita values. This somewhat surprising result may be due to the data quality. Real GDP per capita and real GDP per worker data were obtained from the Penn World Table Mark 5.6 (PWT 5.6). Workforce in PWT 5.6 is defined as the working age population, i.e. the population aged 15-65. Censuses are taken about every 10 years in most countries. Annual population data are obtained by interpolation and workforce data is calculated from these data. Workforce data are probably measured with considerable measurement error. Since the results seem to be robust to the choice of data we decided to concentrate on the per capita variables in order to make our results comparable to two other panel data studies by Islam (1995) and CEL. Some of the results using workforce data are reported in the appendix.

In the Solow model per capita income growth depends on the initial level of per capita GDP, average savings and the average population growth rate (adjusted for the depreciation rate and the rate of technological progress). In the augmented version of the model a measure of schooling is added to the textbook version.

Real GDP per capita, y_{it}, data was obtained from PWT 5.6. RGDPCH is the preferred measure of real per capita GDP, because it is adjusted for purchasing power parity and it is based on a chain index. Quinquennial real per capita GDP data is used, starting in 1960, 1965, ... and ending in 1990.

In our exposition of the Solow model in the previous section we implicitly assumed that investment equals savings. This equality is only true in a closed economy. However, like MRW, Islam (1995) and CEL we proxy the saving rate by the aggregate investment to GDP ratio. The data was obtained from PWT 5.6. The time series were averaged over 1960-64, 1965-69,..., 1985-89.[12]

Population data was obtained from a World Bank source, World Data, available on CD-ROM. It is superior to the population data reported by PWT 5.6, because there are discontinuities for a number of countries in the PWT 5.6 data series around 1970. Only for Taiwan, which is not available from the World Bank, was PWT 5.6 population data used. Average population growth rates were computed as the difference between the natural logarithms of total population at the end and beginning of each period and dividing this difference by the number of years. In this study the average population growth rates for 1960-64, 1965-69, ... , 1985-89 were computed.[13]

Like Islam (1995), MRW and CEL, technological progress and the depreciation rate were assumed to be constant across countries and that they sum to 0.05. The natural logarithm of the sum of population growth and 0.05 was calculated for $ln(n + g + \delta)$.

Instead of proxying human capital investment using school enrolment rates like MRW, we include a measure of the level of human capital in our regressions.[14] Gemmell (1996) shows that proxying the level of human capital by using school enrolment rates is problematic, because it conflates the level and accumulation effects of human capital and leads to misinterpretations

[12] Islam (1995) averages over 1960-65, 1965-70, ... However, averaging over these overlapping periods did not alter our results.

[13] Again, the results remain unchanged when the average rates from 1960-65, 1965-70, ... , 1985-90 are used. Calculating the compound growth rates also did not alter the results.

[14] Please refer to MRW (p 418) for a discussion of including the level of human capital versus the investment in human capital in the Solow model.

of the role of the labour force growth. To avoid these problems we proxy the level of human capital by the average years of schooling.[15] Data collected by Barro and Lee (1996) provides the average years of total schooling for the population aged 15 and older. The data is provided quinquennially, starting in 1960, 1965, ... and ending in 1990. For this study schooling data at the beginning of each five year period was used.

Data is available for 85 countries.[16] However, data is not available for all countries for all of the six periods, thus, making the panel unbalanced.

3.4 Results

3.4.1 The Textbook Solow Model

All reported results in this section are based on the textbook Solow model as shown by equation (3.1). The econometrics package used is DPD98 for Gauss (Arellano and Bond (1998)). All reported standard errors are corrected for heteroskedasticity. The parameter estimates reported from GMM estimation are one-step estimates. All regressions include time dummies which we found to be jointly significant in every regression. In order to conserve space the coefficients on the time dummies are not reported in the tables.

The left hand side variable is the change in the logarithm of real per capita GDP. First, an OLS levels regression was run. All variables are significant at the one percent level and have the expected sign. The negative coefficient on initial GDP as in most published growth regressions is interpreted as conditional convergence while investment is positive and population growth is negative as suggested by the Solow model. The implied speed of convergence, λ, is less than

[15] Bond, Hoeffler and Temple (1998) proxy the level of human capital by the secondary school enrolment rates, but their analysis confirms the main results presented in this chapter.

[16] Please refer to the appendix for a full list of the countries which were included in the sample.

Table 3.1: Textbook Solow Model

	(1) OLS	(2) WG	(3) DIF-GMM	(4) SYS-GMM	(5) DIF-GMM extra instr.	(6) IV Levels
$\ln(y_{t-1})$	-0.035	-0.227	-0.222	-0.100	-0.148	-0.022
	(0.012)	(0.046)	(0.082)	(0.041)	(0.066)	(0.013)
ln(investm.)	0.102	0.159	0.100	0.208	0.157	0.077
	(0.014)	(0.028)	(0.057)	(0.041)	(0.049)	(0.017)
$\ln(n+g+\delta)$	-0.190	-0.381	-0.322	-0.428	-0.350	-0.178
	(0.069)	(0.119)	(0.141)	(0.168)	(0.134)	(0.072)
λ	0.007	0.051	0.050	0.021	0.032	0.004
	(0.002)	(0.012)	(0.012)	(0.007)	(0.011)	(0.003)
m_1	3.48	-4.34	-3.55	-4.36	-3.92	3.34
m_2	1.46	0.20	0.08	-0.05	-0.01	1.40
SarganTest			0.04	0.21	0.12	0.06
Diff. Sargan				0.99	0.82	

Note: Standard errors in parentheses. The test statistics for first and second order correlation are given by m_1 and m_2, respectively (distributed as standard normals). The figures reported for the Sargan tests are the p-values of the null hypothesis, valid specification.

one percent *per annum*. The results are shown in the first column of Table 3.1.

Second a within groups estimator was used, the results are shown in the second column of Table 3.1. Comparing the estimated coefficients of the OLS levels regression and the within groups estimation, we see that the OLS levels regression provides a higher estimate for the coefficient on the lagged dependent variable than the within groups estimation. These estimates provide the approximate upper and lower bounds for the following regressions. Since the coefficient on the lagged dependent variable is lower than the one obtained by OLS levels, the implied speed of convergence based on the results of the within groups estimation is much higher than the estimate obtained from OLS levels, it is about five percent *per annum*.

Column 3 presents the results using the Arellano and Bond (1991) first differenced GMM estimator. We assume that initial GDP is predetermined and investment is endogenous. We

assume that current population growth is exogenous, in the sense of being uncorrelated with shocks to GDP per capita in both the current and preceding five year periods, i.e.

$$E((n+g+\delta)_{it},(v_{it}-v_{i,t-1}))=0$$

This allows the use of both current and lagged levels of population growth as instruments in the first differenced equations[17].

All coefficients are statistically significant, although investment is only significant at the eight percent level. The main problem with these results is that the coefficient on the lagged dependent variable is very close to the within groups estimate. Simulation results reported in Blundell and Bond (1998a) show that the first differenced GMM estimator may be subject to a large downward finite-sample bias in autoregressive models.[18] It may be that the presence of explanatory variables other than the lagged dependent variable, and more particularly the inclusion of current or lagged values of these regressors in the instrument set, will improve the behaviour of the first differenced GMM estimator in particular applications. However, in our case the first differenced GMM estimate seems to be seriously downward biased, because it is very close to the within groups estimate which is expected to be seriously biased downwards in a panel with six or fewer time periods (Nickell (1981)). Since the coefficient estimates obtained from first differenced GMM estimation are close to the ones obtained from within groups estimation the estimated speed of convergence is also similar to the one estimated by

[17] This assumption is weaker than strict exogeneity, since it allows for feedback from earlier GDP shocks onto current population growth. We assess the sensitivity of our results to alternative assumptions about investment and population growth in Tables 3.2 and 3.3 below.

[18] For example, with T=4 and N=100 and a true value of α=0.9, the distribution of the first differenced GMM estimator has a mean of 0.23 (with a standard deviation of 0.83) in Table 2 (a) of Blundell and Bond (1998a). For further details please refer to Table 3.5 in the appendix of this chapter where we present more simulation results.

within groups. The Sargan test only accepts the validity of the instrumental variable set at the four percent level.

CEL also use the Arellano and Bond (1991) first differenced GMM estimator for the same model and report even higher rates of convergence. For the textbook Solow model they report an estimated λ of nearly 13 percent *per annum*. As pointed out in Bond, Hoeffler and Temple (1998) this very high speed of convergence is most probably biased upward. There are two main reasons why CEL obtain such a high λ. First, the coefficient on the lagged dependent variable obtained from differenced GMM is close to or even lower than the coefficient obtained from within groups estimation. Thus, the differenced GMM estimates of the coefficient on the lagged dependent variable appear to be biased downward and as a result the estimated speed of convergence is too high. Second, CEL seem to report the two step GMM estimates of the coefficients, which tend to be lower for the lagged dependent variable than the one step estimator.[19]

We suggest the use of the Blundell and Bond (1998a) system GMM estimator as an alternative to the Arellano and Bond (1991) first differenced GMM estimator. Again treating investment as endogenous and population growth as exogenous in the sense described above, the system GMM results are presented in column 4 of Table 3.1. The coefficient on the lagged dependent variable falls between the 'upper and lower bound', and the estimated speed of convergence is about two percent *per annum*. Since additional moment restrictions are used we expect the system GMM estimation to be more efficient; all standard errors apart from that

[19] Please refer to the appendix section 3.6.3 for a fuller explanation of why we prefer to present the one step results. Using the two step estimates for the differenced GMM regression reported in Table 1 we obtain an implied λ of about six percent *per annum*. The difference between CEL's results and this estimate may be due to the different sample size and time period examined. CEL's study uses data from 97 countries, but only covers 1960-85.

for population growth are smaller than the standard errors obtained from the first differenced GMM estimation. All regressors are statistically significant at the one percent level. For the system GMM estimates the Sargan test does not reject the validity of the instruments, and the Difference Sargan test does not reject the validity of the additional instruments used compared to first differenced GMM.

Thus, system GMM is our preferred estimator for the Solow model for two reasons. The estimate of the coefficient on the lagged dependent variable is not obviously biased, i.e. it lies well above the within groups estimate and well below the OLS levels estimate, and the estimates of the coefficients are more precise than the ones obtained from first differenced GMM. The additional instruments used by the system GMM estimator appear to be both valid and highly informative in this context.

We suspect that the poor performance of the first differenced GMM estimator is due to the weakness of the instrument set used. Blundell and Bond (1998a) argue that the first differenced GMM estimator is likely to perform poorly when the series are close to random walks. Lagged levels are then weak instruments for current differences. As in Bond, Hoeffler and Temple (1998) we consider strengthening the instrument set used by the basic first differenced GMM estimator by including a measure of human capital.

The endogenous growth model as discussed in Romer (1990) provides a theoretical basis for this choice of additional instruments. In this endogenous growth model human capital is a key input in the research sector which generates new designs and these in turn generate new investment opportunities. Based on this assumption we add lagged values of the average years of schooling to the basic instrumental variable set.

The results using this augmented instrumental variable set are shown in column 5. The

Difference Sargan test does not reject the validity of the schooling variable as additional instruments. Notice that, compared to the basic first differenced GMM results, all three coefficients move towards our preferred system GMM estimates. Most importantly, the coefficient on the lagged dependent variable now lies well above the corresponding within groups estimate, and is insignificantly different from the system GMM estimate.[20] Investment, which was only significant at the eight percent level in the basic differenced GMM results, becomes significant at the one percent level when we include our schooling measure as an additional instrument. As we will show in the next section when testing the augmented Solow model, including the average years of schooling as a regressor turns out to be insignificant in most regressions. However, although schooling is insignificant as a regressor, including it in the instrument set appears to strengthen the instrument set significantly.

In column 6 we present the results from the IV levels estimation. The results are similar to OLS levels, and in particular the coefficient on initial income is significantly higher than the system GMM estimate.[21] Both IV levels and OLS levels estimates of this coefficient would be biased upwards by unobserved country specific effects. Thus, we interpret this as evidence for the presence of significant unobserved country specific effects which requires estimation techniques which correctly account for these effects. According to the Sargan test, the validity of the instruments can only be accepted at the six percent level for the IV levels estimator.

Our preferred method of estimation, using the Blundell and Bond (1998a) system estimator,

[20] Using a Hausman test we find that the test statistic is $0.91 \sim N(0,1)$. Thus, we cannot reject the null hypothesis that the coefficients on the lagged dependent variable obtained from the system GMM estimate and the differenced GMM estimate (using additional instruments) are the same.

[21] Using a Hausman test we obtain a test statistic of $2.01 \sim N(0,1)$. Thus, this test rejects the extra moment conditions used by the levels estimator, compared to the system GMM estimator. In other words the null hypothesis that there are no unobserved country specific effects is rejected. Note that this is consistent with the Sargan test for the IV levels estimator.

only provides consistent estimates if there is no serial correlation in the shocks, v_{it} . Thus, we test for second order serial correlation in the first differenced residuals.[22] The hypothesis of no second order serial correlation in the first differenced residuals is not rejected for any of the GMM estimations.

However, for the OLS level estimation as well as for the IV levels estimation we find strong evidence of positive first order serial correlation in the levels residuals, and weaker evidence of higher order serial correlation. We interpret this as further evidence that the error term contains unobserved country specific effects, which causes the levels residuals to be serially correlated.

We now turn to a discussion of investment. In all regressions in Table 3.2 we use system GMM and make the same assumptions about initial per capita GDP and population growth as before, but examine the effect of different assumptions about investment. We can treat investment as either exogenous[23], predetermined or endogenous. One result is common to all regressions presented in this table, no matter which assumption we make about investment, the coefficient on investment is positive and significant at the one percent level. However, the different treatment of investment does change the coefficient on the lagged dependent variable and thus the estimated speed of convergence. If we assume investment to be either exogenous or predetermined the speed of convergence is higher than the two percent estimated from the regression where we assume investment to be endogenous. In either of these alternative specifications, the coefficient on the lagged dependent variable again becomes insignificantly

[22] Please refer to Arellano and Bond (1991). They present a discussion of the test statistic for serial correlation which we report in the tables, it is asymptotically normally distributed. For the OLS levels estimates the serial correlation test tests whether the level of the error terms are uncorrelated, whereas for the GMM estimates, they test whether the first differences of the error terms are uncorrelated.

[23] We assume that current and previous five year period shocks to GDP per capita are uncorrelated with current investment, but that shocks to GDP per capita lagged two or more periods may be correlated with current investment.

Table 3.2: Investment and Growth

	(1) investm.	(2) exog.	(3) investm.	(4) pred.	(5) investm.	(6) endogenous
	full IV set	reduced IV set	full IV set	reduced IV set	full IV set	reduced IV set
$\ln(y_{t-1})$	-0.187	-0.162	-0.217	-0.226	-0.100	-0.095
	(0.042)	(0.041)	(0.039)	(0.040)	(0.041)	(0.043)
ln(investm.)	0.271	0.256	0.317	0.334	0.208	0.207
	(0.032)	(0.032)	(0.035)	(0.039)	(0.041)	(0.044)
$\ln(n+g+\delta)$	-0.501	-0.456	-0.520	-0.520	-0.428	-0.416
	(0.159)	(0.161)	(0.154)	(0.153)	(0.168)	(0.168)
λ	0.041	0.035	0.049	0.051	0.021	0.020
	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)	(0.008)
m_1	-4.42	-4.42	-4.33	-4.28	-4.36	-4.34
m_2	0.09	0.04	0.12	0.12	-0.05	-0.06
SarganTest	0.25	0.06	0.27	0.08	0.21	0.29

Note: Standard errors in parentheses. The test statistics for first and second order correlation are given by m_1 and m_2, respectively (distributed as standard normals). The figures reported for the Sargan tests are the p-values of the null hypothesis, valid specification. We treat y_{t-1} as predetermined and $(n+g+\delta)$ as exogenous.

different from the (biased) within groups estimate.

In this table we also investigate the effects of a reduction in the instrumental variable set. The power of the Sargan test is often low, in particular if the number of instruments used is large in a relatively small sample. For a valid set of instrumental variables we expect that a reduction of the number of instruments used would result in not rejecting the validity of the instrumental variable set at a higher probability level. In column 1 we assume investment to be exogenous, i.e. we assume that both current and previous five year period shocks to GDP per capita are uncorrelated with current investment. In this case we use all current and past levels of investment as instruments in the first differenced equations, and the current first difference of investment as an instrument in the levels equations. According to the Sargan

test the validity of these instruments is not rejected. However, when we reduce the number of instruments set by only using the current and first lagged levels of investment as instruments in the first differenced equations,[24] in column 2, the Sargan test only accepts the validity of the instrumental variable set at the six percent level. Note that the coefficient estimates remain nearly unchanged. Based on the restricted set of instruments, the Sargan test therefore does provide some evidence against the assumption that investment rates can be treated as exogenous. This is consistent with the differences in the estimated coefficients, particularly on the lagged dependent variable and investment, compared to our preferred specification in which investment rates are treated as endogenous.

We observe a similar effect when we assume investment to be predetermined. In this case we assume that current shocks to GDP per capita are uncorrelated with current investment, but that past shocks to GDP per capita may be correlated with current investment. This allows us to use levels of investment lagged one and more periods as instruments in the first differenced equations. The results for the full instrument set, using all available lags, are presented in column 3. Again the Sargan test does not reject the validity of this instrumental variable set. However, when we reduce the instrument set and only use levels of investment lagged once and twice as instruments in the first differenced equations, in column 4, the Sargan test only accepts the validity of the instrumental variable set at the eight percent level.

In the last two columns we treat investment as endogenous (the results for the full instrument set in column 5 are the same as those in Table 3.1 column 4). When we reduce the instrumental variable set here by only using levels of investment lagged two and three times as instruments

[24]Similarly we use only the current and first lag of population growth, and only the second and third lags of GDP per capita - rather than all the available lags - as instruments in these reduced instrument sets.

in the first differenced equations, in column 6, the Sargan test shows no sign of rejecting the validity of this instrumental variable set. Based on these instrument validity tests, and the effect of alternative assumptions on the coefficient estimates, we conclude that the investment rate should indeed be treated as an endogenous variable.

In Table 3.3 we present a similar analysis of alternative assumptions about population growth. We assume here that per capita GDP is predetermined and investment is endogenous. When we treat population growth as predetermined or endogenous, we find similar results to those when we treat population growth as exogenous, albeit with coefficient estimates being less well determined in some cases. The main result here is that these alternative specifications give us no reason to reject the maintained assumption that population growth rates can be treated as exogenous, as in our preferred specification.

3.4.2 The Augmented Solow Model

In this section we present the results obtained from testing the augmented Solow model as described in equation (3.2). The dependent variable is the change in the logarithm of real per capita GDP. It is regressed on the logarithm of initial GDP, the investment to GDP ratio, the population growth rate and the average years of schooling measured at the beginning of the period. Thus, we treat schooling as predetermined.

First, an OLS levels regression was run, the results are reported in column 1 of Table 3.4. As before for the textbook Solow model, the coefficients on all variables are statistically significant and of the expected sign. However, the schooling measure is only significant at the 10 percent level. The estimated speed of convergence, ψ, is about one percent.

Second, a within groups estimator was used, the results are shown in the second column of

Table 3.3: Population Growth

	(1) (n+g+δ)	(2) predetermined	(3) (n+g+δ)	(4) endogenous
	full IV set	reduced IV set	full IV set	reduced IV set
ln(y_{t-1})	-0.081	-0.063	-0.072	-0.067
	(0.044)	(0.049)	(0.042)	(0.044)
ln(investm.)	0.204	0.192	0.195	0.186
	(0.043)	(0.048)	(0.045)	(0.048)
ln(n+g+δ)	-0.258	-0.178	-0.410	-0.436
	(0.187)	(0.208)	(0.231)	(0.264)
λ	0.017	0.013	0.015	0.014
	(0.008)	(0.009)	(0.008)	(0.008)
m_1	-4.40	-4.28	-4.33	-4.32
m_2	-0.13	-0.19	-0.09	-0.09
SarganTest	0.13	0.21	0.08	0.21

Note: Standard errors in parentheses. The test statistics for first and second order correlation are given by m_1 and m_2, respectively (distributed as standard normals). The figures reported for the Sargan tests are the p-values of the null hypothesis, valid specification. We treat y_{t-1} as predetermined and investment as endogenous.

Table 3.4: Augmented Solow Model

	(1) OLS	(2) WG	(3) DIF-GMM	(4) SYS-GMM	(5) IV Levels
$\ln(GDP_{t-1})$	-0.047	-0.229	-0.175	-0.118	-0.040
	(0.014)	(0.046)	(0.071)	(0.052)	(0.015)
ln(investm.)	0.095	0.160	0.152	0.217	0.066
	(0.013)	(0.028)	(0.051)	(0.039)	(0.017)
ln(n+g+δ)	-0.197	-0.367	-0.314	-0.439	-0.188
	(0.068)	(0.128)	(0.146)	(0.163)	(0.071)
ln(schooling)	0.022	-0.017	-0.034	0.005	0.032
	(0.013)	(0.037)	(0.046)	(0.035)	(0.013)
ψ	0.010	0.052	0.038	0.025	0.008
	(0.003)	(0.007)	(0.012)	(0.009)	0.003
m_1	3.53	-7.19	-3.79	-4.40	3.46
m_2	1.43	0.30	0.02	-0.02	1.39
SarganTest			0.12	0.40	0.04
Diff. Sargan				0.98	

Note: Standard errors in parentheses. The test statistics for first and second order correlation are given by m_1 and m_2 , respectively (distributed as standard normals). The figures reported for the Sargan tests are the p-values of the null hypothesis, valid specification.

table 3.4. Schooling is no longer significant. As before the OLS levels regression provides the approximate upper bound and the within groups estimation provides the approximate lower bound for the coefficient on the lagged dependent variable for the following regressions.

Column 3 presents the results using the Arellano and Bond (1991) first differenced GMM estimator, assuming that initial GDP and schooling are predetermined, that investment is endogenous and population growth is exogenous. Here the estimate of the coefficient on the lagged dependent variable falls between the OLS levels and within groups estimates, although it is not significantly different from the within groups estimate. Schooling is again not statistically significant.

Using the system GMM estimator the estimate of the coefficient of the lagged dependent variable falls between the 'upper and lower bounds'. The results are reported in column 4. Schooling is the only variable that is not significant. Again the coefficient estimates are generally more precise than the differenced GMM estimates, and neither the Sargan test nor the Difference Sargan test reject the validity of the additional instruments used.

In the last column of Table 3.4 we present the IV levels estimates.[25] The coefficient on the lagged dependent variable is again close to the OLS levels estimate, and significantly higher than the corresponding system GMM estimate. This is consistent with an upward bias, caused by not taking the unobserved country specific effects into account. Like in the OLS levels results we detect significant first order serial correlation in the levels residuals, which we interpret as further evidence for the presence of unobserved country specific effects. The Sargan test also rejects the validity of the levels IV instrument set at the four percent level.

[25] Notice that in the IV lebels results reported here, we have used a reduced number of instruments, rather than all the available lags. As in Table 3.2, this increased the power of the Sargan test to detect misspecification, but had very little impact on the estimated coefficients.

The coefficient on initial GDP is significantly negative in all five regressions, confirming the so-called 'conditional convergence' hypothesis, a standard result in cross-country growth regressions. Investment is highly significant in all five regressions, even though investment is treated as endogenous in the differenced and system GMM results, as well as in the IV levels estimation. The coefficient on population growth is significantly negative in all regressions. Schooling is statistically significant only in the OLS levels regression and in the IV levels results; it is not statistically significant in any of the results that control for unobserved country specific effects. However, a number of studies have found that the partial correlation between schooling and growth is not robust.[26] The Sargan test of overidentifying restrictions does not reject the instrument set used in the first differenced GMM and system GMM results. Moreover, the Difference Sargan test does not reject the use of the additional instruments in the system GMM estimation in comparison to the differenced GMM estimation. Again the system GMM estimator gives our preferred results. However, it is noticeable that the differences between the first differenced GMM results and the system GMM results are smaller here than they were for the textbook Solow model. Although schooling is not significant once we control for unobserved country specific effects, adding schooling to the instrument set used by the first differenced GMM estimator seems to strengthen the instrument set significantly. Indeed it was this observation that led us to consider using schooling as an additional instrument when estimating the textbook Solow model.

[26] See for example Pritchett (1996a).

3.5 Conclusion

Knowledge of the likely direction of the bias in OLS levels as well as within groups estimation in dynamic models with unobserved country specific effects and a small number of time periods was used to establish approximate upper and lower bounds for the estimated coefficient on the lagged dependent variable in growth regressions. Using differenced GMM estimation we found that the estimated coefficient on the lagged dependent variable appeared to be biased downward. This was particularly the case for the textbook Solow model when no outside instruments were used. Hence, the estimated speed of convergence as reported by CEL on the basis of these differenced GMM estimators is likely to be too high.

We suggest that the differenced GMM estimator is likely to perform poorly when the series are persistent and the number of time series observations is small, because past levels of the series provide weak instruments for the variables in first differences. Estimating the textbook Solow model by differenced GMM, we experimented with strengthening the instrument set by adding a schooling variable to the instrument matrix. Although the coefficient on this schooling variable was found to be insignificant when testing the augmented Solow model, we found that schooling significantly strengthened the basic instrument set for estimating the textbook Solow model in first differences. This gave an estimate of the speed of convergence that was significantly lower than the within groups estimate.

However, a system GMM estimator (using not only lagged levels of the series as instruments in the first differenced equations, but also lagged differences of the series as instruments in levels equations) was shown to be the preferred estimation method, since it produced more precise and more reasonable coefficient estimates than the differenced GMM estimator. The additional

instruments used by the system GMM estimator were not rejected, and were found to be highly informative in these empirical growth models. The estimated speed of convergence based on system GMM is much closer to the conventionally reported two percent *per annum*.

Furthermore, we found that results from the (augmented) Solow model support the conditional convergence hypothesis and investment was found to be highly significant in all regressions, even though investment is treated as an endogenous variable in the preferred system GMM results.

3.6 Appendix

3.6.1 Sample

The sample consists of the following 85 countries :

Benin, Botswana, Cameroon, Central African Republic, Ghana, Kenya, Liberia, Malawi, Mali, Mauritius, Mozambique, Niger, Rwanda, Senegal, Sierra Leone, Sudan, Tanzania, Togo, Uganda, Zaire, Zambia, Zimbabwe, Costa Rica, Dominican Republic, El Salvador, Guatemala, Haiti, Honduras, Jamaica, Mexico, Nicaragua, Panama, Trinidad and Tobago, Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Paraguay, Peru, Uruguay, Venezuela, South Korea, Hong Kong, Singapore, Algeria, Australia, Austria, Bangladesh, Belgium, Burma (Myanmar), Canada, Denmark, Egypt, Finland, France, Germany, Greece, India, Ireland, Israel, Italy, Japan, Jordan, Malaysia, Nepal, Netherlands, New Zealand, Norway, Papua New Guinea, Pakistan, Philippines, Portugal, Sri Lanka, Syria, South Africa, Spain, Sweden, Switzerland, Thailand, Tunisia, Turkey, UK, USA.

3.6.2 Nickell-Bias

To illustrate why the consistency of the within groups estimator depends on large T, we look at a simple autoregressive model with no regressors

$$y_{i,t} = b_0^* y_{i,t-1} + (\eta_i + v_{it}) \text{ for } t = 2, ..., T \tag{3.11}$$

For within groups estimation we transform the model by using the following definitions

$$\overline{y_{it}} = y_{it} - \frac{1}{T-1}\sum_{s=2}^{T} y_{is}$$

$$\overline{y_{i,t-1}} = y_{it} - \frac{1}{T-1}\sum_{s=2}^{T} y_{is}$$

$$\overline{\eta_i} = \eta_i - \frac{1}{T-1}\sum_{s=2}^{T} \eta_i = \eta_i - \eta_i = 0$$

$$\overline{v_{it}} = v_{it} - \frac{1}{T-1}\sum_{s=2}^{T} v_{is}$$

thus, (3.11) becomes

$$\overline{y_{it}} = b_0^* \overline{y_{i,t-1}} + \overline{v_{it}}$$

Thus, this transformation eliminates the country specific effect, η_i. The within estimator is then OLS applied to this transformed equation. $\widehat{b_0^*}$ obtained from within estimation in economic growth panel data with fixed T is biased and inconsistent for the following reason. By definition $y_{i,t-1}$ and $v_{i,t-1}$ are correlated, consequently $y_{i,t-1}$ and $\overline{v_{it}}$ are correlated, since $\overline{v_{it}}$ contains

$v_{i,t-1}$. The correlation between $\overline{y_{i,t-1}}$ and $\overline{v_{it}}$ is negative, thus the estimate of b_0^* will be biased downward. The order of the bias is $O(T^{-1})$,i.e. it gets smaller for larger T.[27]

3.6.3 Arellano and Bond Differenced GMM Estimator

Consider the growth equation (3.6). Taking first differences eliminates η_i and for T time periods we can write [28]

$$\begin{aligned}
\Delta y_{i3} &= b_0^* \Delta y_{i2} + c_0 \Delta x_{i3} + \Delta v_{i3} \\
\Delta y_{i4} &= b_0^* \Delta y_{i3} + c_0 \Delta x_{i2} + \Delta v_{i4} \\
&\vdots \\
\Delta y_{iT} &= b_0^* \Delta y_{i,T-1} + c_0 \Delta x_{iT} + \Delta v_{iT}
\end{aligned}$$

$\triangle y_{i,t-1}$ and $\triangle v_{it}$ are correlated by construction and thus OLS estimates of b_0^* would be inconsistent. However, if the errors terms are not serially correlated, instrumenting for $\triangle y_{i,t-1}$ will provide consistent estimates. In this case y_{it} is a valid instrument in the first period for which we observe the relation, since it is generally correlated with $\triangle y_{i2}$, but not with $\triangle v_{i3}$. In the second period $y_{i1}$and y_{i2} are valid instruments, because $\triangle v_{i4} = (v_{i4} - v_{i3})$ is neither correlated with y_{i1} nor with y_{i2}. Finding instruments for $\triangle y_{i,t-1}$ for all subsequent time periods shows that we are in fact exploiting $M = \frac{(T-1)(T-2)}{2}$ moment restrictions which can be summarized as

$$E(y_{i(t-j)} \triangle v_{i,t}) = 0 \text{ for } j = 2, ..., (t-1) \text{ and } t = 3, ..., T$$

[27] See Nickell (1981) for further details.

[28] This exposition follows Arellano and Bond (1991). Although our panel is unbalanced we concentrate on a balanced panel for expositional purposes.

The next step is to find valid instruments for $\triangle x_{i,t}$. Assuming that $x_{i,t}$ are endogenous and that $x_{i,t}$ is correlated with η_i means that

$$E(x_{i(t-j)} \triangle v_{it}) = 0 \text{ for } j = 2, ..., (t-1) \text{ and } t = 3, ..., T$$

Thus, valid instruments for $\Delta x_{it}$are $(x_{i1}, ..., x_{i,t-2})$.[29] Using the moment restrictions for $y_{i,t-1}$ and x_{it} the variable matrix becomes

$$Z_i = \begin{bmatrix} y_{i1}, x_{i1} & \dots & 0 \\ \vdots & y_{i1}, y_{i2}, x_{i1}, x_{i2} & \vdots \\ & & \ddots \\ 0 & \dots & y_{i1} ... y_{i,T-2}, x_{i,1}, ..., x_{i,T-2} \end{bmatrix}$$

and the moment restrictions can be summarized as

$$E(Z_i^{'} \Delta v_i) = 0$$

where Δv_i is the vector of the differenced error terms $\Delta v_i = (\Delta v_{i3}, ..., \Delta v_{iT})'$. The aim is to obtain an optimal estimator of b_0^* as $N \to \infty$ for fixed T on the basis of these moment restrictions alone. Following Hansen (1982) GMM estimation minimizes the weighted sum of squares $\overline{v}' Z A_N Z' \overline{v}$. The optimal estimator is given by

$$\widehat{\delta} = \begin{bmatrix} \widehat{b_0^*} \\ \widehat{c_0} \end{bmatrix} = \left(W' Z A_N Z' W \right)^{-1} W' Z A_N Z' \Delta y$$

where $W = (\Delta y_{-1}, \Delta X)$, A_N is a square positive definite matrix and ΔX is the $(T-2)N \times K$ stacked matrix of observations on Δx_{it}. Alternative choices of A_N provides one-step or two-step

[29] As Arellano and Bond (1991) note, x_{it} may include a combination of strictly exogenous, predetermined and endogenous variables. For predetermined x_{it} levels of x_{it} lagged one period and more are valid instruments for $\Delta\ x_{it}$.

estimators. The one step estimator is obtained by setting

$$A_N = \left(\frac{1}{N} \sum_{i=1}^{N} Z_i' H Z_i \right)^{-1}$$

where H is a (T-2) square matrix which has twos in the main diagonal, minus ones in the first subdiagonals and zeroes otherwise. This procedure forms the first step of the two step estimator. The residuals obtained from the first step estimation are used to construct a consistent estimate of the variance-covariance matrix of the moment restrictions, $\widehat{V}_N$, i.e. replacing H with $\Delta\widehat{v}_i\Delta\widehat{v}_i'$. In the two step GMM estimation we set $A_N = \widehat{V}_N^{-1}$. The resulting two step estimator is optimal, because the asymptotic variance of the GMM estimator $\widehat{\delta}$

$$\widehat{avar(\delta)} = N \left(W'ZA_NZW \right)^{-1} \left(ZA_NV_NA_NZ'W \right) \left(W'ZA_NZ'W \right)^{-1}$$

is minimized to

$$\widehat{avar(\delta)} = N \left(W'Z\widehat{V_N^{-1}}Z'W \right)^{-1}$$

However, for the special case of spherical disturbances $v_{it} \sim iid(0, \sigma_v^2)$ the one and two step GMM estimator coincide. For more general assumptions about the error terms it can be shown that the one and two step GMM estimators are not asymptotically equivalent. In general the one step estimator will be less efficient.

For this study the one step estimator was the preferred estimator for two reasons. First, Monte Carlo Studies show that the loss in efficiency is not great. Second, the two step GMM estimator converges slowly to its asymptotic distribution. This has implications for hypothesis testing in finite samples. The asymptotic standard errors of the two step GMM are biased downward, which means that the corresponding t-statistics are biased upward. Thus, inference

based on the asymptotic distribution of the two step GMM estimator in finite samples can be seriously misleading.

3.6.4 Blundell and Bond System GMM Estimator

As Blundell and Bond (1998a) show GMM estimators relying on lagged levels as instruments for current differences are likely to perform poorly when the series are close to random walks.[30] They suggest to employ additional moment restrictions in order to be able to use lagged differences as instruments for equations in levels. Recall the simple autoregressive model

$$\begin{aligned} y_{i,t} &= b_0^* y_{i,t-1} + (\eta_i + v_{it}) \text{ or} \\ y_{i,t} &= b_0^* y_{i,t-1} + u_{it} \end{aligned} \tag{3.12}$$

where η_i and $v_{i,t}$ have the familiar error-components structure:

$$E[\eta_i] = 0 \ , \ E[v_{it}] = 0, \ E[v_{it}\eta_i] = 0 \text{ for } i = 1, ..., N \text{ and } t = 2, ..., T \tag{3.13}$$

We assume that the errors are serially uncorrelated

$$E[v_{it}v_{is}] = 0 \text{ for } i = 1, ..., N \text{ and } s \neq t \tag{3.14}$$

and that y_{i1} are predetermined

$$E[y_{i1}v_{it}] = 0 \text{ for } i = 1, ..., N \text{ and } t = 2, ..., T \tag{3.15}$$

[30] This exposition follows Blundell and Bond (1998a).

these assumptions imply the following moment restrictions

$$E\left[y_{i,t-s}\Delta v_{it}\right] = 0 \text{ for } t = 3, ..., T \text{ and } s \geq 2$$

In addition to these previous assumptions we require

$$E\left[\Delta y_{i2} u_{i3}\right] = 0 \text{ for } i = 1, ..., N \tag{3.16}$$

Combining this last assumption with (3.14) and (3.15) implies

$$E\left[\Delta y_{it-1} u_{it}\right] = 0 \text{ for } i = 1, ..., N \text{ and } t = 4, ..., T \tag{3.17}$$

We have now a set of linear moment restrictions implied by (3.15), (3.16) and (3.17). Regarding the additional moment restriction (3.16) we have to examine when it is valid to make this assumption. Given (3.14) and (3.15) the validity of the restriction $E(\Delta y_{i2}(\eta_i + v_{i3})) = 0$ depends on

$$E(\Delta y_{i2}\eta_i) = 0$$

Hence, (3.16) is a restriction on the initial condition process generating y_{i1}. To see this, first write y_{i1} as

$$y_{i1} = \left(\frac{\eta_i}{1 - b_0^*}\right) + e_{i1} \tag{3.18}$$

Now consider equation (3.12) for the first period observed

$$y_{i,2} = b_0^* y_{i1} + (\eta_i + v_{i2})$$

where e_{i1} is the is the initial deviation from the convergent level. Subtracting $y_{i,1}$ from both sides of this equation

$$y_{i2} - y_{i1} = (b_0^* - 1)y_{i1} + (\eta_i + v_{i2}) \tag{3.19}$$

and using (3.18) we obtain

$$\begin{aligned} \Delta y_{i,2} &= (b_0^* - 1)\left(\frac{\eta_i}{1 - b_0^*}\right) + (b_0^* - 1)e_{i1} + (\eta_i + v_{i2}) \\ &= (b_0^* - 1)e_{i1} + v_{i2} \end{aligned}$$

Hence, given assumption (3.14) about the error components structure, assumption (3.16) is equivalent to the restriction $E(e_{i1}\eta_i) = 0$ for $i = 1, ..., N$. A sufficient condition is thus that the initial conditions y_{i1} satisfy the stationarity restriction $E(y_{i1}|\eta_i) = \eta_i/(1 - b_0^*)$ for each individual. Note that this requires only the first moments of the y_{it} series to be stationary, and does not require stationary second moments.

A GMM estimator exploiting these additional moment restrictions requires a stacked system of (T-2) equations in first differences and the (T-2) equations in levels corresponding to periods 3,...,T for which instruments are observed. The instrumental variable matrix is

$$Z_i^+ = \begin{bmatrix} y_{i1} & 0 & \dots & 0 & 0 & 0 & \dots & 0 \\ 0 & y_{i1}, y_{i2} & & & 0 & & & 0 \\ 0 & 0 & \ddots & & \vdots & & & \vdots \\ \vdots & & & y_{i1,\dots,}y_{i,T-2} & 0 & 0 & \dots & 0 \\ - & - & - & - & - & - & - & - \\ 0 & 0 & \dots & 0 & \Delta y_{i2} & 0 & \dots & 0 \\ 0 & & & 0 & 0 & \Delta y_{i3} & & 0 \\ \vdots & & & \vdots & \vdots & & \ddots & \vdots \\ 0 & 0 & \dots & 0 & 0 & 0 & \dots & \Delta y_{i,T-1} \end{bmatrix}$$

Table 3.5: Monte Carlo Simulation Results

N	α	OLS	WG	DIF-GMM	SYS-GMM
100	0.5	0.8736	-0.0370	0.4641	0.5100
		(0.0266)	(0.0697)	(0.2674)	(0.1330)
	0.8	0.9798	0.1343	0.4844	0.8101
		(0.0107)	(0.0726)	(0.8224)	(0.1618)
	0.9	0.9950	0.1906	0.2264	0.9405
		(0.0055)	(0.0725)	(0.8264)	(0.1564)
500	0.5	0.8748	-0.0360	0.4887	0.5021
		(0.0114)	(0.0310)	(0.1172)	(0.0632)
	0.8	0.9801	0.1364	0.7386	0.7939
		(0.0047)	(0.0328)	(0.3085)	(0.0779)
	0.9	0.9951	0.1930	0.5978	0.9043
		(0.0025)	(0.0330)	(0.6407)	(0.0999)

Source: Blundell and Bond (1998a), Table 2(a) for the DIF-GMM and SYS-GMM results. OLS and Within Groups estimates were made available by the authors.

Extending (3.12) by one regressor x_{it} gives

$$\Delta y_{it} = b_0^* \Delta y_{i,t-1} + c_0 \Delta x_{it} + \Delta v_{it}.$$

To exploit the linear moment restrictions in (3.17) and (3.16) we further require that $E(\Delta x_{it}\, \eta_i) = 0$. This allows for the use of lagged Δx_{it} as further instruments in the levels equations.

3.6.5 Comparison of different estimators' finite-sample properties

Table 3.5 reports means (standard deviations) from experiments with T=4 and 1000 replications using Monte Carlo simulations. The model is $y_{it} = \alpha y_{i,t-1} + \eta_i + v_{it}$ with $var(\eta_i) = var(v_{it}) = 1$ and initial conditions drawn from the stationary distribution for y_{i1}. Results are reported for two-step GMM estimators.

3.6.6 Derivation of the speed of convergence

Take the following simple growth model

$$\frac{\ln y(t) - \ln y(0)}{t} \cdot 100 = \gamma \ln y(0) + \beta x(0) + u_{it}$$

the speed of convergence λ can be derived in the following way

$$\lambda = -\frac{\ln(\frac{\gamma \cdot t}{100} + 1)}{t}$$

and its standard error can be computed using the delta method as[31]

$$SE_\lambda = \frac{SE_\gamma}{(\frac{\gamma \cdot t}{100} + 1) \cdot 100}$$

3.6.7 Testing the Augmented Solow Model using Workforce Data

In Table 3.6 we present an estimation of the augmented Solow model. The difference between this table and Table 3.4 is that we use workforce data instead of per capita data. We used GDP per worker data from PWT 5.6 (RGDPW), 'workers' are defined as the population of working age, i.e. aged 15-64. PWT 5.6 does not provide workforce data explicitly. However, the ratio of GDP per capita (RGDPCH) to GDP per worker (RGDPW) gives the proportion of the working age population in the population. We multiplied the population data with this ratio to obtain the workforce data. We then calculated the growth rate of the workforce by taking the log differences between the end and the beginning of each five year period and dividing the difference by five. As in Table 3.4 we assumed the sum of the depreciation rate, δ, and the rate of technological progress, g, to be 0.05.

[31] Please see Goldberger (1991, pp 102-103) for an exposition of the delta method.

Comparing the results from the regressions using workforce data with the ones obtained from the regressions using per capita data we find that the results are very similar. The coefficient on the lagged dependent variable is negative and significant in all regressions, providing further evidence for the conditional convergence hypothesis. Depending on the method of estimation the estimated speed of convergence is between one and six percent *per annum*. For our preferred method of estimation we obtain the same estimated speed of convergence, 2.5 percent *per annum*. Investment is positive and significant in all regressions and the workforce growth rate is negative and significant in all regressions. Our schooling measure is not significant in any regression when we use the workforce data, it is not even significant in the OLS regression (we found schooling to be marginally significant at the ten percent level when we used per capita data). For our preferred method of estimation the Sargan test does not indicate any problems with the validity of the instrumental variable set used. Furthermore, the Difference Sargan test does not reject the validity of the additional instruments in the system GMM estimation in comparison to the first differenced GMM estimation.

Table 3.6: Augmented Solow Model Using Workforce Data

	(1) OLS	(2) WG	(3) DIF-GMM	(4) SYS-GMM
$\ln(\text{GDPW}_{t-1})$	-0.047	-0.260	-0.185	-0.117
	(0.013)	(0.047)	(0.084)	(0.048)
ln(investm.)	0.093	0.163	0.141	0.221
	(0.014)	(0.028)	(0.048)	(0.041)
ln(w+g+δ)	-0.233	-0.368	-0.405	-0.499
	(0.054)	(0.078)	(0.085)	(0.092)
ln(schooling)	0.014	-0.002	-0.000	-0.006
	(0.014)	(0.034)	(0.042)	(0.035)
ψ	0.010	0.060	0.041	0.025
	(0.002)	(0.007)	(0.014)	(0.008)
m_1	3.48	-4.30	-3.73	-4.30
m_2	1.54	0.10	-0.02	-0.07
SarganTest			0.08	0.35
Diff. Sargan				0.99

Note: Standard errors in parentheses. The test statistics for first and second order correlation are given by m_1 and m_2 , respectively (distributed as standard normals). The figures reported for the Sargan tests are the p-values of the null hypothesis, valid specification.

Chapter 4

Openness, Investment and Growth

4.1 Introduction

In the previous chapter we used two relatively simple models, the Solow growth model and its augmented version, for our empirical analysis of economic growth in a cross-section of countries. In this chapter we want to examine a more general model, following Sachs and Warner (1997), who augment the neo-classical growth model by adding a number of policy and structural variables. Lack of openness to international trade and structural characteristics, such as being landlocked and having a tropical climate, are identified as important determinants of economic growth. Another interesting result is that the investment to GDP ratio is insignificant in all of their regressions. This confirms the findings of an earlier paper (Sachs and Warner (1995)) which examines the effects of openness and investment on growth in the original Barro regression. Controlling for other variables they find that open countries had higher growth rates. Investment on the other hand is not significant at the five per cent level in their specification. Analyzing the relationship between investment and openness in a regression of investment on openness and initial GDP they find that initial GDP and openness are both positive and sig-

nificant. Richer and more open economies have higher investment. Sachs and Warner conclude that poor trade policies reduce investment as well as growth. This suggests that openness is already helping to account for the difference in investment ratios across countries, thus, causing the coefficient on investment to be insignificant.[1]

The insignificance of investment in their growth regressions is a surprising result in the light of the majority of empirical growth studies and the previous chapter which find the investment ratio to be highly significant. Levine and Renelt (1992) as well as Sala-i-Martin (1997a, 1997b) test for the robustness of investment in cross country growth regressions and find strong evidence for a robust positive link between growth and investment in single cross-country growth regressions. Both studies use Leamer's extreme bounds analysis to identify 'robust' empirical relationships in growth regressions. Sala-i-Martin's (1997a) conclusions are based on the results of approximately 4 million regressions. He examines two different base models, the first includes initial per capita income, initial primary enrolment rates and initial life expectancy as the 'fixed' variables, the second model adds the average investment to GDP ratio to the set of 'fixed' variables. He then randomly adds three further regressors from a pool of 58 variables which have been used in cross country growth regressions and examines the entire distribution of the coefficients on each regressor. If the fraction of the cumulative distribution function which lies above (or below) zero is 95 per cent or greater, Sala-i-Martin concludes that this variable is robustly correlated with growth.[2] Investment and the Sachs and Warner measure of openness are among the variables which are 'robustly' correlated with growth. However, there is no discussion of the relationship between investment and openness

[1] Sachs and Warner (1995, pp 47-52) and Sachs and Warner (1997, p 351).
[2] For a more detailed discussion please refer to section 2.4.1.

in Sala-i-Martin (1997a).

The aim of this chapter is to re-examine the Sachs-Warner hypothesis that openness is such an important determinant of investment that investment is no longer significant in growth regressions when we control for openness. We focus on two possible explanations of why investment may not be significant in a single cross-country growth regression. First, openness and/or other regressors may be highly correlated with investment, thus, causing the estimated coefficient on investment to be insignificant. We will therefore test whether the result is robust to different specifications. Second, ignoring unobserved country specific effects and endogeneity problems may cause biased parameter estimates in which investment appears to be insignificant. We will therefore examine whether our preferred treatment of the dynamic relationships in a panel data framework produces a different result.

This chapter is structured in the following way. In the second section we present the basic model and discuss the Sachs and Warner (1997) single cross-country results. Section three presents the empirical results from our single cross-country regressions and from the panel estimation. The last section concludes.

4.2 The Basic Model

Sachs and Warner (1997) test a model in which they regress growth on a number of variables, including a number of time invariant country characteristics. We discuss the data in more detail in the next section, here we concentrate on the basic model and the main findings.

In the basic model the average annual growth rate of per capita GDP from 1965 to 1990 is regressed on initial per worker GDP, a measure of average openness throughout the period, an interaction term between openness and initial GDP, a landlocked dummy, a dummy for tropical

climate, the log of life expectancy and its square, average government saving as a proportion of GDP, a measure of the quality of public sector institutions, the ratio of primary commodity exports to GDP, and the difference in the growth rates of the workforce and the population. The analysis is based on a single cross-country regression.

Using the Sachs and Warner (1997) data set we replicate exactly their basic regression in the first column of Table 4.1.[3] The main results are that the coefficient on openness is positive and significant and that the coefficient on the interaction term is negative and significant. The interpretation of these findings is that openness affects growth in two ways. A higher degree of openness is associated with higher per capita GDP in the steady state and with a faster rate of convergence to the steady state. The time invariant country characteristics, landlocked and tropical climate, have significantly negative coefficients. Being landlocked acts as a natural barrier to trade which will reduce growth. Having a tropical climate reduces growth through a number of channels, e.g. harvests are less reliable and tropical diseases reduce the productivity of the work force.

Life expectancy, government savings and the quality of institutions have positive and significant coefficients. Higher life expectancy proxies a healthier and more productive workforce.[4] Growth is also fostered by higher government savings, because they tend to raise the national savings rate and thus promote capital accumulation which is often believed to be beneficial for growth.[5] Higher institutional quality means that property rights are well defined and enforce-

[3] We would like to thank the authors for the use of their data set. These are the same results as in Sachs and Warner (1997) their Table 2, column 1. Variable definitions are given in the data appendix.

[4] However, there are two alternative explanations of the relationship between life expectancy and income, as for example Pritchett and Summers (1996) point out. The first possibility is that higher income causes higher life expectancy (reverse causation). The second possible explanation is that one or several factors may cause both higher life expectancy and higher income (incidental association).

[5] For a critical discussion of this hypothesis please refer to King and Levine (1994).

Table 4.1: The Sachs-Warner Growth Model

n = 79	(1)	(2)	(3)	(4)	(5)
$\ln Y_{65}$	-1.632	-1.631	-1.200	-1.225	-1.886
	(0.193)	(0.194)	(0.177)	(0.206)	(0.193)
$\ln Y_{65}$*open	-0.766	-0.765	-0.733	-0.355	0.262
	(0.301)	(0.304)	(0.338)	(0.331)	(0.036)
open	8.475	8.457	8.311	5.400	
	(2.462)	(2.488)	(2.782)	(2.817)	
landlocked	-0.580	-0.581	-0.658	-0.860	-0.697
	(0.215)	(0.217)	(0.244)	(0.281)	(0.231)
tropics	-0.849	-0.845	-0.935	-0.776	-0.838
	(0.233)	(0.240)	(0.269)	(0.302)	(0.258)
ln Life Exp.	45.473	44.885			51.820
	(17.481)	(18.693)			(19.990)
(ln Life Exp.)^2	-5.395	-5.323			-6.239
	(2.244)	(2.387)			(2.552)
cgovsav	0.123	0.122	0.107		0.132
	(0.023)	(0.023)	(0.026)		(0.249)
institutions	0.283	0.281	0.179		0.180
	(0.072)	(0.075)	(0.075)		(0.074)
sxp	-3.265	-3.276	-3.636	-3.956	-2.810
	(0.957)	(0.972)	(1.075)	(1.240)	(1.035)
gw-gpop	1.192	1.189	1.749	1.956	1.568
	0.313	(0.317)	(0.327)	(0.375)	(0.319)
ln investm.		0.020	0.438	0.656	0.078
		(0.213)	(0.215)	(0.246)	(0.228)
λ	0.021	0.021	0.014	0.015	0.026
	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)
adj. R^2	0.89	0.89	0.86	0.83	0.87

Note: The dependent variable is the annual average growth rate 1965-90. Standard errors in parentheses. Sachs-Warner (1997) data used. All specifications include an intercept term.

able and hence growth rates tend to be higher in countries with strong institutions.

A high ratio of primary commodity exports to GDP is negative for growth. This is interpreted as evidence for the hypothesis that the industrial structure of an economy is important for growth. More endogenous growth effects are possible in the tradeable manufacturing industry rather than in the primary commodity tradeables sector. Dynamic Dutch disease effects can also be at work, reducing growth through a high primary commodity export share in GDP.

Implicitly this empirical specification is based on a neo-classical growth model. The Solow model examines per worker income growth which depends on technological progress, the saving ratio and the growth rate of the labour force. However, in their specification Sachs and Warner chose to regress the per capita GDP growth rate on initial per worker GDP and the difference between the growth rate of the workforce and the population. As expected, the coefficient on the difference between the two growth rates is positive, indicating that growth of per capita GDP tends to be higher in countries in which the workforce grows at a higher rate than the population.

In their study Sachs and Warner include a discussion of the omitted variable problem. In the neo-classical growth model the omitted variable is the initial level of technology, A(0).[6] Initial income and initial technology will tend to be positively correlated, thus causing an upward bias of the coefficient on initial income if the initial level of technology varies across countries. The authors therefore suggest including the most comprehensive possible list of observable indicators of A(0) in the specification. They claim that, since their regressions explain nearly 90 per cent of the variation in growth across countries, the size of this bias is probably small.[7] However, in

[6] For a more detailed discussion of the initial level of technology please refer to Chapter 3.

[7] The extent of an omitted variable bias cannot be assessed by the size of R^2. In general, R^2 measures the explained variation in the regression. There are different sources of variation in a multiple regression, variation unique to each regressor and variation common to two or more regressors (unless they are uncorrelated). If initial

a single cross-section regression we cannot assess how large the bias is. Instead we propose to examine the extent of this bias in a panel regression, because in the previous chapter we found the biases from both country effects and simultaneity to be important.

There are several possible sources for simultaneity biases in this type of regression. We are particularly concerned about the inclusion of life expectancy as a regressor. While a healthier and longer living population is likely to be more productive, i.e. average income is likely to be higher, income is also likely to be an important determinant of life expectancy. Most obviously at the microeconomic level, higher income results in better nutrition and thus better resistance against illness. Richer households can also spend more on health care.[8] At the macroeconomic level, wealthier countries are able to spend more resources on public health programmes which reduce mortality rates.

According to a detailed World Bank (1984) study, the higher a country's income the higher is life expectancy. This is also confirmed by our descriptive statistics in Table 4.2, the correlation coefficient of income per worker and life expectancy is 0.85. Higher incomes are associated with lower mortality rates and since life expectancy at birth measures the number of years a newborn baby could be expected to live if his or her mortality pattern exactly paralleled that of all age groups in the current year, higher income is also associated with higher life expectancy.

This positive association between income per capita and life expectancy is depicted in Figure 4.1. We fitted a log linear regression line, low per capita income countries had low average life expectancy and high per capita income countries had a higher average life expectancy, but the increase in life expectancy is not proportionate. This may be due to the fact that very

income is highly correlated with the omitted variable we could in principle obtain a high R^2, while the coefficient on initial income remains severely biased due precisely to this correlation.

[8]For a detailed microeconomic theoretical framework, please refer to Behrman and Deolalikar (1988).

high income countries have low child mortality rates, thus the average life expectancy is more determined by adult mortality. However, causes of child and adult mortality differ.[9] The main causes of child mortality are communicable diseases while the main causes of adult deaths are heart and lung disorders, which are related to the consumption of tobacco and alcohol, which tends to rise with income.

While the World Bank (1984) study emphasises the important role of public health programmes, in particular vaccination and antimalarial spraying, there are other important determinants of life expectancy. Improved infrastructure, in particular better sanitation and access to safe water, as well as (female) education are important determinants of life expectancy. It was found that female education reduces infant mortality significantly, for example in Latin America infant mortality rates were five times higher among children whose mothers have no schooling than among those with mothers with 10 or more years of schooling.[10]

Thus, it is likely that unobserved country specific effects, such as infrastructure and education, are common determinants of income per capita and life expectancy.

We suggest that because of a possible simultaneity bias, life expectancy should not be treated as an exogenous variable in growth regressions, particularly when unobserved country specific effects are not controlled for. More generally, not accounting for unobserved country specific effects is likely to cause an omitted variable bias. We examine these problems using a panel data approach in Section 4.3.3.

An interesting result of the Sachs and Warner (1997) study is that contrary to the predictions of the Solow model, the average investment ratio is insignificant in their specification. We

[9] See Pritchett and Summers (1996).
[10] World Bank (1984, p 70).

confirm this finding by including the natural logarithm of the investment to GDP ratio in their model. The results are shown in column 2 of Table 4.1. The coefficient on investment is insignificant and the other coefficients remain nearly unchanged.[11]

Investment may be insignificant, because openness and/or other regressors are highly correlated with investment. In this case investment might only explain a small amount of variation in growth independently of openness (or other regressors). Openness could then be significant while investment is not, due to multicollinearity. Sachs and Warner (1995, 1997) suggest that poor trade policies reduce investment as well as growth. Thus, openness would already account for the differences in investment ratios across countries, causing the coefficient on investment to be insignificant. However, we can show that it is the inclusion of life expectancy, and not openness, that accounts for the insignificance of investment in the Sachs-Warner model.

Table 4.2 presents the correlation coefficients between the policy variables, life expectancy and investment. Notice that whilst there is a positive correlation between openness and investment, there is an even higher correlation between life expectancy and investment. Indeed, apart from government saving, the correlation between life expectancy and all of the other variables is strong and positive, between 0.69 and 0.85. Thus, including life expectancy as a regressor in a growth model is likely to reduce the significance of some of the other variables.

Openness as measured in this study proxies openness to trade and capital flows. It is also highly correlated with most other variables other than government saving. The more open a country is the more investment takes place, mainly because many countries import a large amount of investment goods. A better institutional framework is also associated with higher

[11] It is not quite clear in Sachs and Warner (1997) whether they include the level or the log of investment in their regression. We tried both, level and log, and found that the results are very similar. Throughout this chapter we report the regressions including the natural logarithm of investment.

Table 4.2: Correlation Coefficients

	$\ln Y_{65}$	ln investment	openness	ln Life Exp.$_{70}$	cgovsav
ln Y_{65}	1				
ln investment	0.59	1			
openness	0.59	0.63	1		
ln Life Exp.$_{70}$	0.85	0.73	0.69	1	
cgovsav	-0.13	0.03	-0.13	-0.12	1
institutions	0.72	0.62	0.74	0.71	-0.06

Note: Data from Sachs and Warner (1997) used.

investment.

Life expectancy and investment are highly correlated, probably because they are both individually highly correlated with initial GDP per worker. As discussed above, higher income is associated with higher life expectancy and as discussed in Chapter 3 investment is an important determinant of per worker income. According to the data government saving and total investment are uncorrelated.

We examine the possibility of multicollinearity problems by dropping the variables suspected of causing the problem. The key finding is that once we omit the life expectancy terms the point estimate of investment increases while its standard error remains nearly unchanged.[12] Investment is now significant at the five percent level. The coefficients on all the other regressors are not greatly affected. In particular, both investment and the openness terms are significant in this specification. These results are presented in column 3 of Table 4.1.

If we further reduce the model by dropping the institutional quality variable and the government savings rate the coefficient on investment becomes significant at the one percent level.

[12] Most econometric textbooks concentrate on high standard errors as a symptom of multicollinearity. However, as Maddala (1992, Chapter 7) shows high intercorrelations among regressors do not necessarily result in high standard errors, because the standard error of the coefficient depends not only on the correlation among the regressors, but also on the estimated variance of the error term and the variation in the regressors.

These results are shown in column 4. Note that although the other coefficients remain very similar in comparison with the model presented in column 3, the interaction term between initial GDP per capita and openness is no longer significant in this reduced model.

We also examine the possibility that it may be the inclusion of the openness measure which causes investment to be insignificant. However, when we drop openness from the Sachs and Warner model investment remains insignificant. These results are shown in column 5 in Table 4.1.[13]

To summarize, we suggest that because of a possible simultaneity bias, life expectancy should not be treated as an exogenous variable in growth regressions, particularly when unobserved country specific effects are not controlled for. Testing different specifications of the Sachs-Warner growth model we showed that the coefficient on investment is positive and significant once we drop the life expectancy terms. Openness to international trade as well as the investment ratio are positive and significant in this reduced specification. Thus, it is not openness which causes investment to be insignificant, but life expectancy.

4.3 Re-examination of the Basic Model

The aim of this section is to re-examine the basic model. We want to examine the various issues discussed above, in particular the insignificance of investment, the omitted variable and simultaneity problems, in a dynamic panel data framework applying the methods described in Chapter 3. However, as we will discuss in more detail below, it is not straightforward to construct a panel data set which corresponds to the Sachs-Warner single cross-country data set

[13]The coefficient on investment becomes marginally significant (p=0.12) when we drop openness as well as the interaction term of openness and initial GDP per worker.

for two reasons. First, some of the data series from the original data sources quoted by Sachs and Warner were slightly different from the data used in their study. Second, it turned out to be impossible to obtain time varying measures for all of the variables. Thus, we decided to construct two data sets, a single cross-country data set and a five year panel data set using data from the sources quoted by Sachs and Warner. We proceed by describing the data used for both data sets. Then we try to reproduce the main features of the Sachs-Warner results using our single cross-country data. Finally, we use our five year panel data to re-estimate the basic model.

4.3.1 Data

In order to construct our single cross-country data set and our panel data set we divide the regressors into three categories. These are genuinely time invariant country characteristics, time varying variables for which we have complete time series data and time varying variables for which we have little or no time series data.[14] Measuring the values of the first two categories should prove relatively straightforward. For time varying variables five-year averages or five-yearly observations can be obtained from the original data sources and for the genuinely time invariant variables we use the values as provided by Sachs and Warner. These time varying variables are per worker GDP, openness, the share of primary commodity exports in GDP and the difference in the growth rates of the workforce and the population. For all of the descriptive statistics please refer to the appendix.

The log of the initial real GDP per head of the economically active population was obtained from PWT 5.6. The economically active population is defined as the population aged 15-64.

[14] A more detailed discussion of the data is presented in the appendix.

Our measure of GDP per head of the economically active population as measured in 1965 differs slightly from the measure used by Sachs and Warner. The descriptive statistics are presented in the appendix. Unfortunately, we were unable to establish the cause of this difference.

The annual values of the openness measure were obtained from Sachs and Warner (1995). This binary measure of openness is constructed in the following way. A country is classified as closed if it has at least one of the following characteristics (1) nontariff barriers covering 40 percent or more of the trade, (2) average tariff rates of 40 percent or more, (3) a black market exchange rate that is depreciated by 20 percent or more relative to the official exchange rate (4) a socialist economic system, and (5) a state monopoly on major exports. The openness dummy takes a value of zero when the country is closed and a value of one when the country is open. The openness measure used in the single cross country regression is the average annual value over the period 1965-90.

The share of primary commodity exports in GDP can be obtained from the World Bank. However, Sachs and Warner (1997) improve the data quality and quantity by using alternative sources[15]. We only used the data as provided by the World Bank World Data CD Rom 1994.

As in the previous chapters the investment to GDP ratio was obtained from PWT 5.6.

For the genuinely time invariant country characteristics, the landlocked and tropical climate dummies, we use the values provided by Sachs and Warner (1997).

Unfortunately, the remaining regressors in the Sachs-Warner model are variables which can vary over time, but for which there are little or no time series data available and which are therefore more difficult to include in our panel data set. These variables are life expectancy, government saving and the institutional quality index.

[15] For further details please see the data appendix in Sachs and Warner (1997).

Sachs and Warner state in their data appendix that they obtained their measure of life expectancy from Jong-Wha Lee, but no detailed reference is given. We compared their life expectancy data with data from the World Development Indicators (WDI) 1997.

For 1970 we found that the WDI and the Sachs-Warner data were quite different. For the WDI data the mean is 58.3 years, but the mean of the Sachs and Warner data is 55.3 years. Thus, on average life expectancy is three years lower in the Sachs-Warner data set. Since we were not able to obtain similar life expectancy data, we decided to use the Sachs-Warner measure of life expectancy in the single cross country regressions. In the appendix we report the results of the regressions using the WDI life expectancy data. As could be expected with such differences in the data, the coefficient estimates differ. However, as we will show the main results are robust to using either measure.

For the panel estimation we obtained time varying life expectancy data by imputing quinquennial data from the available WDI data.[16] Using this data, we find that our results are robust to alternative assumptions about the exogeneity in the panel growth regression.

Government saving is defined as current revenues minus current expenditure as a percentage of GDP. Some time varying data is available from the WDI, however for many (developing) countries this data is measured too infrequently in order to predict missing values.

The institutional index is a composite index of institutional quality measures as used by Knack and Keefer (1995). It is an unweighted average of five indices measuring different aspects of institutional quality, such as the rule of law, bureaucratic quality, corruption in government, risk of expropriation, and government repudiation of contracts. 1982 is the earliest year for

[16] Life expectancy data is available from the WDI for 1970, 1972, 1977, 1980, 1987 and 1992 for most countries. To impute missing values for 1965, 1975 and 1985, we fitted a linear trend for each country through these observations.

which the data is available, thus, we do not have annual values for the entire period 1965-90.

Since life expectancy, government saving and institutional quality are only measured infrequently or the data series are too short, we decided to treat them initially as time invariant variables in the panel. However, given the importance of the life expectancy variable in the previous section, we also present panel results using the available time varying data on life expectancy.

To summarize, we use the same data as Sachs and Warner (1997) for the genuinely time invariant country characteristics and for the assumed time invariant variables, and data from the original data sources for the remaining time varying variables. We would therefore expect to be able to obtain similar results to Sachs and Warner (1997) when the same estimation method is used. We now turn to the discussion of the empirical results.

4.3.2 Single Cross-Country Estimation

Using our single cross-country data set we estimate the Sachs-Warner model by OLS to see whether we can reproduce the results obtained in section 4.2. We present the results using our cross-country data set in Table 4.3.

Comparing these results with those obtained using the Sachs-Warner data set we find that the point estimates and standard errors are broadly similar for all variables apart from the primary commodity export share. The much reduced point estimate of the coefficient on the primary commodity export share causes the coefficient to become insignificant. The data we use is the World Bank data, whereas Sachs and Warner took great care to improve the World Bank data. However, although our primary commodity export share is insignificant, it does not cause any other important changes in the results in the single cross-country regression. Since

Table 4.3: Single Cross-Country Growth Regressions

n = 79	(1)	(2)	(3)	(4)
$\ln Y_{65}$	-1.324	-1.328	-0.936	-1.004
	(0.202)	(0.204)	(0.182)	(0.207)
$\ln Y_{65}$*open	-0.780	-0.781	-0.819	-0.675
	(0.344)	(0.346)	(0.371)	(0.360)
open	9.113	9.131	9.566	8.447
	(2.943)	(2.966)	(3.193)	(3.204)
landlocked	-0.641	-0.638	-0.716	-0.989
	(0.261)	(0.263)	(0.286)	(0.321)
tropics	-1.131	-1.143	-1.130	-1.084
	(0.299)	(0.308)	(0.335)	(0.377)
ln Life Exp.	41.544	43.120		
	(20.957)	(22.543)		
(ln Life Exp.)^2	-4.899	-5.092		
	(2.687)	(2.875)		
cgovsav	0.149	0.150	0.138	
	(0.027)	(0.028)	(0.030)	
institutions	0.161	0.166	0.110	
	(0.083)	(0.086)	(0.084)	
sxp	-0.898	-0.869	-1.357	-0.393
	(0.988)	(1.005)	(1.082)	(1.222)
gw-gpop	1.461	1.469	1.921	1.967
	0.359	(0.364)	(0.369)	(0.424)
ln investm.		-0.051	0.380	0.606
		(0.257)	(0.247)	(0.280)
λ	0.016	0.016	0.011	0.012
	(0.003)	(0.003)	(0.002)	(0.003)
adj. R^2	0.83	0.84	0.81	0.74

Note: The dependent variable is the annual average growth rate 1965-90. Standard errors in parentheses. Sachs and Warner (1997) data used for landlocked, tropics, life expectancy, institutions and government savings. All other variables were obtained from the original data sources. All specifications include an intercept term.

the only major difference between their cross-section results and ours is the significance of the primary commodity export share, we put this difference down to the difference in the data for this variable and not to any methodological differences.

In Table 4.3 column 2 we see that as before investment is insignificant when we add it to the Sachs-Warner growth model. When we drop the life expectancy terms in column 3, the coefficient on investment again increases substantially, and becomes significant at the 13 percent level. When we reduce the model further by dropping the institutional quality variable and government savings, investment becomes significant at the five percent level.

To summarize, the previous estimation results from different specifications of the Sachs-Warner growth model are broadly confirmed when we use the data from the original data sources. The estimated speed of convergence is as previously estimated between one and two percent *per annum*. One important difference is that the primary commodity export share is not significant. Nevertheless, the coefficients on openness and investment are both positive and significant in a model which does not include life expectancy terms, the institutional quality index and government saving.

The Sachs-Warner model is likely to suffer not only from multicollinearity problems, but also from omitted variable bias and endogeneity problems. Their model does not allow for unobserved country specific effects, such as for example technology and infrastructure, which we found to be significant in the previous chapter. Furthermore, there are no clear guidelines in the literature on how to treat life expectancy. It is as likely to be a result of income as a cause of income.[17] This endogeneity problem is not addressed in the single cross-country regression

[17] Using panel data Easterly (1997) examines income as a determinant of quality of life. When he regresses life expectancy on income he finds that the coefficient on income is positive and significant in an OLS levels regression.

analysis. In order to address the omitted variable problem as well as the endogeneity issues we turn now to a dynamic panel data estimation of the Sachs-Warner model.

4.3.3 Panel Data Estimation

Estimation Method

In this section we want to use panel data to investigate the effects of openness and investment on growth when we control for unobserved country specific effects. In order to investigate these problems we divide the entire period 1965-90 into five-year subperiods. We can sketch the Sachs-Warner model as follows

$$gy_{it} = \beta x_{it} + \gamma w_i + \eta_i + v_{it} \tag{4.1}$$

where t denotes points in time $t = 2, ..., T$. Here, gy_{it} denotes growth over a five year period, which starts in $t-1$ and ends in t. The first time period for which the growth rate is observed is 1965-70, the second period 1970-75, ... , and the last period 1985-90. The x_{it} are either period averages or the initial value taken at the beginning of each period in 1965, 1970, ..., 1985. For example openness and investment are average values whereas initial GDP per worker and the primary commodity export share are measured at the beginning of each period. The w_i denote the measured time invariant country characteristics, tropical climate and landlocked, and the quasi-fixed variables, life expectancy, government savings and institutional quality, which are here assumed to be the same for all subperiods. Since we found unobserved country specific effects to be significant in the previous chapter we also allow for unobserved country specific effects, η_i. Finally, v_{it} denotes an error term which is assumed to be serially uncorrelated.

Since the measured country characteristics w_i may be correlated with the unobserved coun-

try specific effects and/or the error term v_{it} we estimate the model in the following two step procedure. First, we estimate (4.1) without including the measured country specific characteristics, w_i

$$gy_{it} = \beta x_{it} + \eta_i^* + v_{it} \tag{4.2}$$

where $\eta_i^* = \gamma w_i + \eta_i$. This allows us to obtain consistent estimates of the β coefficients, whatever the correlation between w_i and the error components. Notice that the main coefficients of interest here, on investment and openness, can be estimated in this way. In order to obtain a consistent estimate of $\widehat{\beta}$ we use the Blundell and Bond (1998a) system GMM estimator.[18] We then use this consistent estimate $\widehat{\beta}$ to estimate the residuals of equation (4.2). In the second step we regress these residuals on the measured country characteristics, w_i

$$(gy_{it} - \widehat{\beta}\, x_{it}) = \gamma w_i + (\eta_i + v_{it}) \tag{4.3}$$

OLS levels estimates of (4.3) will generate consistent estimates of γ if and only if all w_i characteristics are uncorrelated with η_i. Since this is a very strong restriction, not much causal significance should be attached to the coefficient estimate $\widehat{\gamma}$. In general, this two step procedure will not solve the problem of biased estimates of $\widehat{\gamma}$. We present these results for descriptive reasons only, because they describe which measured country characteristics are correlated with variations in the unobserved country specific effects. The advantage of this two step procedure, compared with estimating equation (4.1) by system GMM under the assumption that the measured country characteristics, w_i, are all uncorrelated with the unobserved country specific

[18] For a detailed discussion of this estimator please refer to Chapter 3.

effects, η_i, is that the latter procedure would also give biased and inconsistent estimates of the β coefficients if this very strong assumption is not correct. Due to the presence of the unobserved country specific effects, we also expect the error term $(\eta_i + v_{it})$ of equation (4.3) to exhibit positive serial correlation.

Results

We start by estimating equation (4.2), the 'step one regression', as outlined above. We estimate the basic Sachs-Warner growth model without the time invariant country characteristics landlocked and tropics as well as the quasi-fixed variables life expectancy, institutional quality and government saving, using the system GMM estimator. In comparison to the single cross-country regressions, the sample size is reduced from 79 countries to 76.[19] We treat initial GDP per worker and primary commodity exports as predetermined, because they are measured at the beginning of each period. The difference between the workforce and population growth rates as well as the investment share were treated as endogenous. We only include openness and not the interaction term between openness and initial income per worker, because we found that when we included openness and the interaction term in the same panel regression, both coefficients were individually insignificant. Since our results with respect to investment were not sensitive to whether we included the interaction term or not, we report results for the specification that includes only the level of the openness variable. However, we retained lagged values of the interaction term in the instrumental variable set.

The system GMM results for this version of the Sachs-Warner growth model are presented in column 1 of Table 4.4. The coefficient on the initial GDP per worker is negative and significant

[19] There was insufficient data to include Belgium, Taiwan and Jordan in the panel.

and the estimated speed of convergence is 3.4 percent *per annum*. Both coefficients on openness and investment are highly significant and positive. The coefficient on the difference between the workforce and population growth rates is no longer statistically significant, suggesting either that this variable may be highly correlated with the unobserved country specific effects, or that there is insufficient variation over time to identify the effect of the growth rate differential once we control for unobserved country specific effects. The positive and statistically significant coefficient on the primary commodity export share is surprising. In the single cross-country regressions using the same measure, the coefficient on the primary commodity export share was insignificant, perhaps this result was due to not controlling for unobserved country specific effects. However, the quality of the primary commodity export share data used in this section is questionable and thus all findings with respect to the explanatory power of the primary commodity export share should be treated with caution. The Sargan test does not indicate any problems with the validity of the instrumental variable set used. Finally, we cannot reject the hypothesis of no second order serial correlation in the first differenced residuals.

We also estimated the same model by OLS levels, the results are presented in column 2. A comparison of the OLS levels estimates with the system GMM confirms that the estimated coefficient on initial GDP per worker is higher in the OLS levels regression, confirming our findings from the previous chapter. The OLS coefficient is likely to be upward biased, because the estimation does not allow for unobserved country specific effects. Therefore, the estimated speed of convergence is lower than that estimated from the system GMM results. The estimated speed of convergence is about one percent which is very similar to estimates obtained from the single cross-country regression presented in Table 4.3. Furthermore, like in the single cross-country regressions, the coefficient on the difference between the workforce and population

Table 4.4: Panel Data Estimation

n = 367 76 countries	(1) SYS-GMM	(2) OLS levels	(3) SYS-GMM Life Exp. endog.	(4) SYS-GMM Life Exp. exog.
$\ln Y_{i,t-1}$	-3.085	-0.912	-3.108	-3.749
	(0.905)	(0.175)	(1.111)	(1.275)
open	3.168	2.251	1.698	2.101
	(1.079)	(0.352)	(0.806)	(0.674)
ln Life Exp.			10.581	10.536
			(6.412)	(7.449)
sxp	11.320	0.232	11.381	11.877
	(4.913)	(1.762)	(4.722)	(4.837)
gw-gpop	0.844	1.247	0.293	0.461
	(0.856)	(0.378)	(0.908)	(1.025)
ln investm.	4.022	1.333	3.069	3.425
	(0.733)	(0.281)	(0.795)	(0.898)
λ	0.034	0.009	0.034	0.042
	(0.011)	(0.002)	(0.013)	(0.016)
m_1	-4.26	1.66	-4.34	-4.24
m_2	-0.67	0.93	-0.45	-0.45
Sargan Test	0.18		0.20	0.13

Note: The dependent variable is the annual average growth rate 1965-90. Standard errors in parentheses. Period dummies are included in each of these specifications. The test statistics for first and second order correlation are given by m_1 and m_2, respectively (distributed as standard normals). The figures reported for the Sargan test are the p-values of the null hypothesis, valid specification.

growth rates is positive and significant while the coefficient on the primary commodity export share is not significant. For the OLS levels regression we find positive first order serial correlation in the levels residuals. The serial correlation between the error terms is positive, thus providing additional support to our assumption that the error term contains unobserved country specific effects.

Estimating this version of the Sachs-Warner model by OLS levels rather than by system GMM shows that the coefficients on openness and investment are still positive and significant. The estimated long run elasticity of per capita GDP with respect to the investment share is 1.3 for the system GMM results in column 1, and 1.5 for the OLS levels results in column 2. For comparison, the long run elasticity suggested by the OLS estimates of the cross-section growth regressions in column 4 of Table 4.3 is about 0.5.[20] We have presented evidence to suggest that both OLS estimates of this long run elasticity are potentially biased, and if this is the case then there is no reason why they should be biased in the same way. For example, a simultaneity bias caused by correlation between *transient* shocks to per capita GDP growth and current investment would have a bigger impact on the panel OLS estimates than on the cross section OLS estimates, since the latter are biased on average values over a longer period of time.[21] This may explain why the panel OLS estimate of this elasticity is higher than the cross section OLS estimate.

We now turn to the 'step two regression', in which we use the coefficient estimates from the system GMM regression presented in column (1) of Table 4.4 to construct the levels residuals, which we regress on the time invariant country characteristics landlocked and tropics and on

[20] Evaluated at the mean value of openness.

[21] For example, a positive transient shock associated with high investment in one five year period, followed by a negative transient shock associated with low investment in the next five year period, would have no impact if we use data for 10 year (or longer) averages.

the quasi-fixed variables life expectancy, institutions and government savings. The results are presented in the first column of Table 4.5. We find that the coefficients on the genuinely time invariant country characteristics, landlocked and tropics, are both negative and significant, as in the single cross-country regression results. The institutional quality index, which was measured in 1982, is found to be negative and significant at the five percent level, and neither government savings nor life expectancy are significant.[22] However, when we regress the residuals on the log of life expectancy alone we find the coefficient on life expectancy to be positive and significant (column 2). Thus, we do find some evidence of a positive correlation between life expectancy and the unobserved country specific effect, η_i, although this relationship is dominated by the inclusion of the landlocked and tropical climate variables. Life expectancy is negatively correlated with both landlocked ($\rho = -0.26$) and with tropical climate ($\rho = -0.66$). Thus, when we include landlocked and tropical climate in the 'step two regression' we find that the coefficient on life expectancy becomes insignificant.

As suggested in our discussion of the estimation method above, we find the residuals to be highly serially correlated in both these regressions, this is consistent with the presence of unobserved country specific effects, η_i. Moreover, as noted earlier, we should not attach causal significance to these coefficient estimates, because these estimates are consistent if and only if all the measured country specific characteristics and the unobserved country specific effects are uncorrelated, which is a strong restriction.

Our main finding from Sections 4.2 and 4.3.2, that both openness and investment have significant positive effects in the growth model, are thus confirmed by our panel data analysis. We

[22] When we included the square of the log of life expectancy we found that the coefficients on the two life expectancy terms were also jointly insignificant.

Table 4.5: Measured Country Characteristics

	(1)	(2)
landlocked	-1.849	
	(0.850)	
tropics	-3.050	
	(0.554)	
ln life exp.	0.898	3.485
	(2.461)	(1.494)
cgovsav	0.027	
	(0.084)	
institutions	-0.336	
	(0.1667)	
m_1	3.34	3.71
m_2	2.53	3.07
adj. R^2	0.18	0.05
F-test	p=0.000	

Note: Standard errors in parentheses. Life expectancy data from Sachs and Warner (1997) used. Both specifications include an intercept term. The test statistics for first and second order serial correlation are given by m_1 and m_2, respectively (distributed as standard normals).

check the robustness of these results by including our time varying measure of life expectancy in the 'step one regression'. We report results only for specifications with a linear life expectancy term, because we found that when we included both life expectancy and its square the two coefficients were estimated very imprecisely. Although there is considerable variation in life expectancy across countries, it varies little over time. This may be the reason why we cannot identify separate effects from life expectancy and its square when we control for unobserved country specific effects. Treating life expectancy as endogenous and correlated with the unobserved country specific effects, η_i, in column 3 we find that the coefficient on life expectancy is positive and significant at the ten percent level. The coefficients on initial GDP per worker and the primary commodity exports remain nearly unchanged, while the coefficients on openness and investment are slightly reduced, but both are still positive and significant. The difference in the workforce and population growth rate is insignificant as before.

We carry out a second robustness check, presented in column 4, where we treat life expectancy as exogenous and uncorrelated with the unobserved country specific effects, η_i. Although life expectancy itself is not statistically significant here, the coefficients on the other variables are again similar to those presented in columns (1) and (3).

To summarize, we find allowing for both unobserved country specific effects and the endogeneity of the regressors, the coefficient on investment is positive and significant in a regression which accounts for openness to international trade. This result is robust to the inclusion of a time varying life expectancy term, irrespective whether we treat life expectancy as endogenous or as exogenous. We find that once we control for unobserved country specific effects, η_i, the significance of life expectancy is greatly reduced in comparison to the single cross-section results.

4.4 Conclusion

In this chapter we examined Sachs' and Warner's (1995, 1997) suggestion that openness to international trade is such an important determinant of investment that if a growth regression already controls for openness then investment will be insignificant, while the coefficient and the standard error of openness will remain nearly unchanged. This result is confirmed when we include investment in the single cross-section regression of the Sachs-Warner model. However, we find that this result is not robust to different specifications of their model. Excluding life expectancy is sufficient to obtain a positive effect of investment, that is significant at the five per cent level. When we exclude life expectancy and the institutional quality index, investment is significant at the one per cent level. Thus, we find that investment is significant in a growth regression controlling for openness. In the Sachs-Warner specification it is mainly the life expectancy variable, and not openness, which reduces the significance of investment.

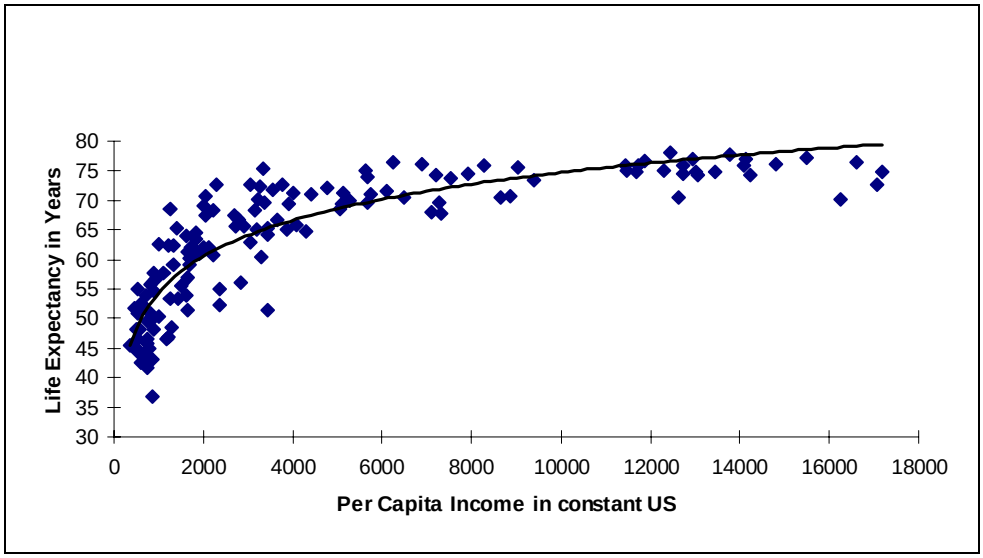
However, growth regressions that include life expectancy as an explanatory variable are likely to suffer from endogeneity problems, because per capita income is likely to be a cause as well as the result of life expectancy. Moreover, there may be omitted variables, such as sanitation and education that affect both life expectancy and income. Thus, we suggest that life expectancy should not be included in growth regressions without addressing this simultaneity problem.

As suggested in Chapter 3, single cross-country growth regressions are likely to produce biased estimates of the coefficient on initial income due to an omitted variable problem. In neo-classical growth models we must account for the differences in initial technology across countries. Sachs and Warner suggest that controlling for observed country characteristics such as climate and landlocked, and policy differences such as openness to trade, may be sufficient to

eliminate most of the bias caused by unobserved country specific effects. However, we find that unobserved country specific effects are significant and ignoring these effects introduces a serious omitted variable bias. When we account for unobserved country specific effects and address the endogeneity issues using a system GMM estimator, we again find that investment is significant in the growth regressions despite controlling for openness. In the panel data analysis, this result is found however we treat life expectancy. We also found that once we control for unobserved country specific effects the significance of life expectancy is greatly reduced, whether we treat life expectancy as a fixed or time varying characteristic. To sum up, we confirm that the Sachs-Warner measure of openness has a significant, positive effect on income in their sample of countries. However, this variable is not so important that it drives investment out of the model. This appears to be an artefact of other problems with the specification and estimation method used in the Sachs-Warner analysis.

4.5 Figures

Figure 4.1: Per Capita Income and Life Expectancy, 1987



4.6 Appendix

4.6.1 List of Variables

lny$_{t-1}$is the logarithm of the initial real GDP per head of the economically active population. GDP data are from the Penn World Table 5.6, RGDPW series were used.

gy is the average annual growth rate of real GDP per capita, data are from Penn World Table 5.6, RGDPCH series were used. Denoting the length of the time period with t the average annual growth rate was computed in the following way:

$$gy = \frac{\ln y(t) - \ln y(0)}{t} * 100$$

open is the fraction of years in which the country is rated as an open economy according to Sachs and Warner (1995).

sxp is the initial share of primary commodity exports in GDP. Data was obtained from the World Bank World Data CD Rom 1994.

gw-gpop is the difference between the average annual growth rate of the workforce and the population. It was computed from the workforce and population data as provided by the World Bank World Data CD Rom 1994.

tropics is the approximate fraction of a country's land area that is subject to a tropical climate. Source: Sachs and Warner (1997).

landlocked is a dummy variable which takes the value 1 if a country is completely landlocked and 0 otherwise. Source: Sachs and Warner (1997).

ln life exp is the logarithm of life expectancy measured around 1970. Source: Sachs and Warner (1997). The time varying measure of life expectancy was obtained from the World

Bank World Development Indicators (WDI). Life expectancy data is available from the WDI for 1970, 1972, 1977, 1980, 1987 and 1992 for most countries. To impute missing values for 1965, 1975 and 1985, we fitted a linear trend for each country through these observations.

cgovsav is the average value of central government savings over the period 1970-90. Savings is defined as current revenues minus current expenditure and is measured in per cent of GDP. Source: Sachs and Warner (1997).

institutions is a measure of institutional quality. It is an unweighted average of five indices measuring different aspects of institutional quality, such as the rule of law, bureaucratic quality, corruption in government, risk of expropriation, and government repudiation of contracts. The data used by Sachs and Warner were measured in 1982. The original data was collected by Political Risk Services and published in the International Country Risk Guide. The individual variables are either on a scale of 0-6 or 0-10. A high value of this index means that the quality of the institutions is high. For further details please refer to Knack and Keefer (1995).

ln investment is the natural logarithm of the ratio of real domestic investment (private plus public) to real GDP. Data source is Penn World Table 5.6.

4.6.2 Descriptive Statistics

Table 4.6 presents the means and the standard deviations for the single cross section data sets. The left part of the table describes the data as used by Sachs and Warner (1997) and the right part describes the data as used in our single cross section data set which was used for the regressions in Table 4.3.

As can be seen from Table 4.6 the means and standard deviation of the variables in both data sets are the same or they are very similar. However, for some unknown reason the real

GDP per head of the economically active population from PWT 5.6 differs slightly from the series used by Sachs and Warner. PWT 5.6 provides a number of different real GDP series and we examined the per capita, per worker and per equivalent adult series. None provided exactly the same descriptive statistics as the Sachs and Warner (1997) data. In their paper they refer to the GDP series as the real GDP per head of the economically active population which is defined as the population between the ages 15-64. This description is the same as used in PWT 5.6 for the per worker GDP (RGDPW series). Thus, we decided to use the per worker GDP data. The GDP per capita data are also not exactly the same, thus the average growth rates differ slightly, too.

We obtained the life expectancy as of 1970 data from the World Bank WDI. Data is not provided for Taiwan and Jordan. For these countries we assumed the same value as given by Sachs and Warner. We found that our life expectancy measure was on average 3 years higher than the measure used by Sachs and Warner. The difference between the life expectancy for China between the two sources was very high (about 15 years), thus, we decided to use the Sachs and Warner value for China.

For the primary commodity variable our values differ slightly from the Sachs and Warner measure. However, we only used the data as provided by the World Bank World Data CD-ROM 1994 and did not attempt to improve the data as suggested and carried out by Sachs and Warner (1995).

Sachs and Warner obtained the difference in the workforce and population growth rates from the World Bank Data CD-ROM 1995. Unfortunately, this data source was not available to us and we used the World Bank World Data CD-ROM 1994. This may account for the slight difference in the data.

Table 4.6: Single Cross-Country Data - Descriptive Statistics

	Sachs and Warner Data		1965-1990 Data	
n = 79	Mean	Std. Dev.	Mean	Std. Dev.
gy	1.903	1.889	1.893	1.886
$\ln y_{t-1}$	8.192	0.884	8.535	0.991
open	0.4153	0.449	0.4153	0.449
ln Life Exp	3.987	0.227	4.041	0.210
sxp	0.115	0.092	0.120	0.110
gw-gpop	0.248	0.336	0.232	0.334
ln investm.			2.713	0.604

Table 4.7: Time-Invariant and Quasi-Fixed Variables - Descriptive Statistics

n=79	Mean	Std. Dev.
tropics	0.546	0.483
landlocked	0.165	0.373
cgovsav	1.557	3.413
institutions	5.93	2.354

Table 4.7 presents some descriptive statistics for the time invariant country characteristics and the quasi fixed variables.

In Table 4.8 we show the means and standard deviations for the data used in the panel data set. This data set only includes 76 countries, because there was no data available for the primary commodity export share for Jordan, Taiwan and Belgium.

Table 4.8: Panel Data - Descriptive Statistics

n = 367, 76 countries	Mean	Std. Dev.
gy	1.852	3.045
$\ln Y_{t-1}$	8.821	0.997
open	0.398	0.477
sxp	0.136	0.118
gw-gpop	0.244	0.509
ln investm.	2.698	0.641

4.6.3 Additional Regression Results

In the following table we present the estimation of the same specifications as in Table 4.3 columns 1 and 2. The only difference is that we use life expectancy data from WDI. The two life expectancy terms are not individually significant, but are jointly significant at the two percent level. As before the primary commodity export share is not significant. The institutional quality index is only marginally significant at the 12 percent level.

Table 4.9: Single Cross-Country Growth Regressions

n = 79	(1)	(2)
$\ln Y_{65}$	-1.270	-1.262
	(0.218)	(0.221)
$\ln Y_{65}$*open	-0.761	-0.760
	(0.354)	(0.357)
open	9.005	8.974
	(3.045)	(3.068)
landlocked	-0.628	-0.633
	(0.273)	(0.275)
tropics	-1.178	-1.160
	(0.029)	(0.322)
ln Life Exp	33.403	31.353
	(24.701)	(25.994)
(ln Life Exp)^2	-3.854	-3.606
	(3.140)	(3.291)
cgovsav	0.156	0.154
	(0.312)	(0.030)
institutions	0.145	0.140
	(0.088)	(0.090)
sxp	-1.020	-1.057
	(1.032)	(1.047)
gw-gpop	1.557	1.543
	(0.373)	(0.378)
ln investm.		0.017
		(0.263)
λ	0.015	0.015
	(0.003)	(0.003)
adj. R^2	0.83	0.82

Note: Standard Errors in parentheses. Dependent Variable is the average annual growth rate 1965-90. Life expectancy data from WDI 1997 used. Both specifications include an intercept term.

Chapter 5

A Re-examination of the Relationship between Growth and some Country Characteristics

5.1 Introduction

This chapter presents a re-examination of the relationship between growth and some country characteristics. A number of growth studies find country characteristics, such as location and cultural factors, to be significant in growth regressions. For example considerable attention has been paid to the slow growth performance of Sub-Saharan Africa. Among others Barro (1991), Levine and Renelt (1992) and Sala-i-Martin (1997a, 1997b) find that the coefficient on a dummy variable for African countries is negative and significant in a number of different specifications. Similarly, the same studies find that a Latin America dummy is negative and significant, and for example Barro (1997) finds that an East Asia dummy is positive and significant. The significance of the coefficients of these regional dummies is interpreted as evidence that some regularities are missing from the model, i.e. that these growth models can neither fully account

for Africa's and Latin America's low growth performance nor for the East Asian economies' high growth performance.

In addition, the importance of cultural factors for economic growth has been highlighted. Barro (1997) and Sala-i-Martin (1997a, 1997b) include the proportion of the population affiliated with major religions and find that five religious variables are robustly correlated with growth: Eastern Religions, Buddhist and Muslim (positive) and Protestant and Catholic (negative). Easterly and Levine (1997, 1998) find that ethnic diversity is negatively correlated with growth.

None of these studies have controlled for unobserved country specific effects or for the endogeneity of investment ratios, which we have found to be important in the previous chapters. Using the same 'two step estimation' procedure as in Chapter 4[1], we therefore re-examine the significance of these measured country characteristics. We use the preferred coefficient estimates of the augmented Solow model obtained by system GMM estimation in Chapter 3 as the 'step one' regression results and calculate the residuals. In the 'step two regression' we regress these residuals on the measured country characteristics. We estimate this 'step two regression' by OLS, correcting the standard errors for serial correlation and for heteroskedasticity. We want to stress, however, that the coefficient estimates on the measured country characteristics are consistent if and only if all the measured country characteristics are uncorrelated with the unobserved country specific effects. Since this is a strong assumption, we do not want to attach any causal significance to these coefficient estimates. The 'step two regression' results describe which measured country characteristics can account for some of the variation in countries' growth performance that is not explained by the augmented Solow model.

[1] *cf.* section 4.3.3.

This chapter is structured in the following way. In the second section we examine the relationship between location and growth. In section three we concentrate on the analysis of religious variables in growth regressions. This is followed by a brief analysis of ethnic diversity and growth. The last section concludes.

5.2 Location

A number of growth studies, for example Barro (1991, 1997), Levine and Renelt (1992) and Sala-i-Martin (1997a, 1997b), examine whether certain regions experienced significantly different growth rates from the rest of the world. These studies include regional dummies in their single cross-country growth regressions, which take a value of one if the country is located in the particular region and zero otherwise. In a Barro growth regression, a significant coefficient on a regional dummy means that this region's growth performance on average was significantly different from the rest of the World's during the sample period, after accounting for the effects of the included regressors. Thus, a significant coefficient on a regional dummy indicates that some regularities are missing from the model to explain the difference in regional growth, i.e. these regressions are not able to explain *why* these regions' growth rates have been different. Common findings are that countries in Sub-Saharan Africa and Latin America experienced on average lower growth rates, while East Asian countries experienced higher growth rates.

A number of papers try to establish why regional growth rates differed. Roubini and Sala-i-Martin (1992) for example present evidence that financial repression may account for Latin America's growth performance. When they include a measure of financial repression in the original Barro regression, the coefficient on the Latin America dummy becomes insignificant. Stokey (1994) suggests that the difference in female and male education can account for much

of the high growth rates experienced by East Asian economies.

Africa's low growth performance and possible explanations have attracted considerable attention in the literature.[2] There are two major studies in which the Africa dummy is not significant. In Easterly's and Levine's (1997, 1998) core growth regression the Africa dummy is negative and significant, however adding the average growth rate of the country's neighbours (weighted by their GDP) makes the Africa dummy insignificant. Furthermore, the study finds that neighbourhood spillovers magnify the effects from bad policies, such as high government deficits, repressed financial markets, distorted exchange rates and political instability. However, why these neighbourhood spillovers matter is not clear. The study suggests that since adaptation of technology to a local environment is risky, a foreign direct investor may find it easier to invest in a neighbouring country if the investment was successful in the first country. A neighbours' success may also have demonstration effects and reputation effects. However, trade does not appear to be the channel for spillover effects. Using trade weights Easterly and Levine found the spillover variable to be insignificant. Although the Africa dummy is insignificant in their study we remain sceptical as to how far the inclusion of a neighbourhood spillover effect really explains Africa's growth performance, because a neighbourhood variable is very similar to the concept of a regional dummy, i.e. for each African country, the neighbours will be other slow growing African countries.

Sachs and Warner (1997) pursue a different approach in their examination of the African

[2]One potentially important question is whether the conventionally used empirical growth models are really adequate to explain Africa's growth, or whether Africa is just too different from other regions and thus Africa specific models should be analysed. Savvides (1995), Ojo and Oshikoya (1995) as well as Hadijmichael et al. (1995) use commonly used specifications of growth models in order to examine Africa's growth performance. Rather than using a global sample they only include African countries in their studies. Their analyses show that the commonly found results, for example conditional convergence and a strong positive effect of investment on growth, hold for this reduced sample. We interpret this as evidence that there is no need to concentrate on an Africa specific model.

growth performance. They specify essentially a neo-classical model and add their openness to international trade measure and a number of geographical variables, such as a landlocked dummy and a tropical climate dummy. The model is a single cross-country growth regression and the estimation method is OLS. An Africa dummy is added to the core model and found to be insignificant. Thus, Sachs and Warner suggest that poor policies, most importantly lack of openness to international trade, as well as geographical factors such as tropical climate and being landlocked are the key variables in explaining Africa's poor growth performance. However, we will show that once we allow for unobserved country specific effects in estimation, the Africa dummy is found to be insignificant even in a basic augmented Solow growth model.

Before turning to our regression results we present some descriptive statistics for our panel data set by region. The sample is exactly the same as in Chapter 3, it consists of 85 countries, and a maximum of six five-year periods per country (1960-90). The panel is unbalanced and the list of countries by region is presented in the appendix. As can be seen from Table 5.1 Sub-Saharan Africa (SSA) was the region which experienced by far the lowest growth rates *per annum* (0.52 percent). Latin American economies grew on average by 1.17 percent, also well below the sample average of 1.89 percent.

Turning to the explanatory variables we see that the average initial income per capita for Sub-Saharan African countries was very low, only about 37 percent of the average initial income per capita. Thus, according to the conditional convergence hypothesis African countries should have experienced higher than average growth rates, *ceteris paribus*. However, on average Sub-Saharan African countries invested less of their GDP (9.02 percent) than the average country (16.29 percent), they had higher population growth (2.21 percent) than the average country (1.66 percent) and much lower initial schooling (1.27 years) than the average country (3.47

years). These differences should account for some of Africa's low growth performance.

As our descriptive statistics for Latin America show, all explanatory variables' means are relatively close to the sample means. Initial income (2318 US$) was higher than the average income (2250 US$) which may account for some of the lower growth rates, *ceteris paribus*. Latin America's initial schooling (3.33 years) and average investment (14.97 percent) were slightly below average (3.47 years and 16.29 percent respectively) and population growth (1.88 percent) was slightly above average (1.66 percent). Again, we expect that these differences should be able to account for some of Latin America's low growth performance.

The three Asian Tiger economies included in our sample, South Korea, Singapore and Hong Kong,[3] grew on average more than three times faster than the average country. Below average initial income per capita (1603 US$), higher than average investment to GDP ratio (23.54 percent), higher initial schooling (4.58 years) and lower than average population growth (1.60 percent) may account for some of the growth rate differential.

Table 5.1: Descriptive Statistics

	Sample	OECD	non-OECD	SSA	Latin America	Asian Tigers
per capita GDP growth*	1.89	2.93	1.52	0.52	1.17	6.49
real GDP per capita (US$)**	2250	5592	1487	835	2318	1603
investment to GDP ratio*	16.29	25.87	12.95	9.04	14.97	23.54
population growth*	1.66	0.61	2.03	2.21	1.88	1.60
initial total years of schooling**	3.47	6.56	2.39	1.27	3.33	4.58
number of countries	85	22	63	22	21	3

*Notes: * indicates annual averages given in percent, ** measured for the earliest available year, for most countries in 1960*

We now turn to our regression results. First, we include dummies for African, Latin Amer-

[3]There was insufficient data to include Taiwan, the fourth so called Tiger economy.

ican and the Asian Tiger economies in the augmented Solow model. The dummies take the value of one if the country is located in the region, and zero otherwise. We first estimate this model by OLS levels, the results are presented in the first column of Table 5.2.

The coefficients on the Africa and Latin America dummies are found to be significant and negative, while the coefficient on the Asian Tiger dummy is found to be positive and significant. This confirms the commonly found results in the literature when unobserved country specific effects and endogeneity are not accounted for, i.e. the Solow growth model appears to be unable to account for the growth experience of African and Latin American and Asian Tiger economies.

Note however that we find significant positive autocorrelation in the residuals from the OLS regression. This is consistent with the presence of unobserved country specific effects which are not accounted for in the OLS regression.

Since the OLS estimates suffer from fixed effects and endogeneity biases we now use our preferred 'two step estimation' procedure to estimate the coefficients on the regional dummies. We use the consistent coefficient estimates from system GMM estimation of the augmented Solow model to calculate the residuals of this growth regression. For convenience we repeat the system GMM coefficient estimates in Table 5.2, column 2.[4] Notice that the effects of investment and population growth are found to be substantially higher in these system GMM results. We then regress these residuals on the regional dummy variables, the results are presented in column 3. We find that the coefficients on the Africa and Latin America dummies are insignificant, but the coefficient on the Asian Tiger dummy is positive and significant. These results are confirmed when we only include one regional dummy at a time in our regression (columns 4, 5 and 6).

[4] These are the same estimates as in Table 4.4, column 4.

Table 5.2: Regional Differences in Growth

	(1) OLS	(2) SYS-GMM	(3) Step 2	(4) Step 2	(5) Step 2	(6) Step 2
lny_{t-1}	-0.052	-0.118				
	(0.010)	(0.052)				
$ln(inv)$	0.078	0.217				
	(0.012)	(0.039)				
$ln(n+g+\delta)$	-0.156	-0.439				
	(0.054)	(0.163)				
$ln(schooling)$	0.016	0.005				
	(0.011)	(0.035)				
SSA	-0.074		-0.017	-0.019		
	(0.025)		(0.031)	(0.030)		
$Latin$	-0.057		-0.018		-0.020	
	(0.014)		(0.018)		(0.019)	
$ATiger$	0.171		0.153			0.162
	(0.015)		(0.036)			(0.036)
m_1	2.11	-4.40	4.50	4.80	4.72	4.46
m_2	-0.05	-0.02	3.45	3.70	3.59	3.35
adj R^2	0.285		0.042	0.001	0.001	0.044
F-test	p=0.000		p=0.000			

Note: Standard errors in parentheses. n=489, 85 countries

Neither the Africa dummy nor the Latin America dummy are required to account for the variations in growth rates once we control for the presence of unobserved country specific effects in estimating the model. This suggests that the augmented Solow model can account for the difference in the growth performance experienced by the African and Latin American countries. Given our preferred coefficient estimates, the low investment ratios and high population growth rates of the African and Latin American countries are sufficient to explain their slow growth of income per capita. However, there is a positive and significant association between the growth regression residual and the Asian Tiger dummy. Thus, the augmented Solow model does not seem to be able to fully account for the high growth rates experienced by the Asian Tiger economies.

5.3 Religion

Sala-i-Martin (1997b) lists five religious variables which he found to be robustly correlated with growth: Eastern Religions, Buddhist and Muslim (positive) and Protestant and Catholic (negative). These variables give the fraction of a country's population affiliated with major religions measured in 1970. The source for this data is Barro (1997) who also presents a detailed discussion of the data[5]. Catholic includes Greek and Russian Orthodox, while Protestant includes Anglican and other marginal Protestant groups, such as Jehovah's Witnesses, Mormons and new age cults. Eastern Religions include Confucianism, Chinese folk religions, Shinto and new religionists, but excludes Buddhists.

We added these five variables to the augmented Solow model and first estimated this model

[5]We would like to thank Robert Barro for the use of his data set. The original data was collected for the World Christian Encyclopedia (Barrett (1982)). Religious affiliation is determined by a person's answer to the question 'What is your religion?' and not by church attendance or taking part in other religious activities.

Table 5.3: Religion and Growth

	(1) OLS	(2) step 2	(3) step2
$\ln y_{t-1}$	-0.041		
	(0.111)		
ln(inv)	0.084		
	(0.012)		
ln(n+g+δ)	-0.268		
	(0.060)		
ln(schooling)	0.027		
	(0.013)		
Catholic	-0.024	-0.013	
	(0.316)	(0.049)	
Protestant	-0.071	-0.058	-0.065
	(0.036)	(0.057)	(0.026)
Buddhist	0.038	0.002	
	(0.043)	(0.061)	
Eastern Rel.	0.339	0.304	0.310
	(0.053)	(0.094)	(0.068)
Muslim	0.041	0.040	
	(0.038)	(0.060)	
m_1	1.97	4.47	4.48
m_2	-0.36	3.16	3.24
adj R^2	0.26	0.06	0.06
F-test	p=0.000	p=0.000	p=0.000

Note: Standard errors in parentheses.
n=489, 85 countries

by OLS levels. The results are presented in Table 5.3. According to the OLS levels estimation presented in column 1 only Protestant and Eastern Religions are significant at the ten and one percent level respectively. When we run our preferred 'step two regression', regressing the growth regression residuals on these five religious variables, we find that only Eastern Religions is statistically significant. However, when we drop Catholic, Buddhist and Muslim from the specification, Protestant and Eastern Religions are both significant at the one percent level. The coefficient on Protestant is negative, whereas the coefficient on Eastern Religions is positive.

Table 5.4: Eastern Religions

	Eastern Religions (%)	Growth Rates (%)
Hong Kong	54.8	6.29
Singapore	54.7	6.65
South Korea	25.7	6.66
Malaysia	25.4	4.28
Japan	24.5	5.26
Thailand	1.8	4.45

The economies with the highest fraction of people affiliated with Eastern Religions in our sample are the three Asian Tiger economies as shown in Table 5.4. Thus, the significant positive coefficient on the Eastern Religions variable may simply indicate the regional differences in growth. In order to examine this possibility we regress the growth regression residuals on the five religious variables as well as the three regional dummy variables. The results from our preferred two step estimation procedure are presented in Table 5.5. In this general specification only the coefficient on the Asian Tiger dummy is significant. The coefficient on Eastern Religions is positive, but not significant at conventional levels (p=0.15). However, although none of the religious variables are individually significant, they are jointly highly significant according to a Wald test (p=0.001). A Wald test for joint significance of the regional dummies also finds them to be jointly significant (p=0.051). In the second column we only include Protestant, Eastern Religions and the Asian Tiger dummy in the regression. In this parsimonious specification all three variables are individually significant at the one percent level.

All six countries with a high proportion of the population affiliated with Eastern Religions (as listed in Table 5.4) are amongst the top ten growth economies. Thus, the Eastern Religions variable accounts for some of high growth performance of Japan, Malaysia and Thailand, and not only for the growth performance of the three Asian Tigers. However, the significance of

Table 5.5: Regional and Religious Differences in Growth

	(1) step 2	(2) step 2
Catholic	-0.031	
	(0.057)	
Protestant	-0.082	-0.066
	(0.062)	(0.026)
Buddhist	-0.020	
	(0.067)	
Eastern Rel.	0.129	0.192
	(0.089)	(0.064)
Muslim	0.025	
	(0.061)	
SSA	-0.028	
	(0.034)	
Latin	-0.010	
	(0.022)	
ATiger	0.077	0.068
	(0.031)	(0.018)
m_1	4.42	4.46
m_2	3.17	3.21
adj. R^2	0.08	0.06
F test	p=0.000	p=0.000

Note: Standard errors in parentheses.
n=489, 85 countries

the Asian Tiger dummy indicates that the Eastern Religions variable cannot fully account for the high growth performance of Hong Kong, South Korea and Singapore. Thus, these results suggest that the augmented Solow model cannot fully account for the high growth performance of these countries even when we allow for differences in religion.

We now turn to a brief discussion of the Protestant religion variable. Table 5.6 lists the ten countries with the highest proportion of Protestants. The average growth rate of these countries was 2.60 percent, which is above the sample average of 1.93 percent. Thus, although their growth rates were above average, the explanatory variables of the Solow model predict

Table 5.6: Protestant Religion

	Protestant Religion (%)	Growth Rates (%)
Norway	98.5	3.26
Denmark	96.0	2.41
Finland	95.0	3.26
New Zealand	76.0	1.23
UK	75.0	2.20
Jamaica	74.7	1.20
Sweden	72.9	2.22
Papua New Guinea	62.0	0.48
Australia	61.4	2.06
USA	52.4	2.00

even higher growth rates for these countries. With the exception of Jamaica and Papua New Guinea all of the countries with a high proportion of Protestants are OECD countries. The Protestant affiliation variable may reflect regional effects such as being a Scandinavian country or a former British colony. The growth rates of the former British colonies appear to have been low over this period. These countries may have certain institutional characteristics in common which are suspected of lowering growth rates. Barro (1997) suggests that Protestant countries tend to have a higher level of democracy and high levels of democracy are also associated with lower growth rates. However, we leave the examination of the relationship between political variables and growth for future research.

5.4 Ethnic Diversity

Easterly and Levine (1997, 1998) find that the partial correlation between ethnic diversity and growth is negative. They argue that higher ethnic diversity is associated with high political instability and bad public policies, such as for example low schooling, underdeveloped financial systems and insufficient infrastructure, and may be related to civil conflicts, as we discuss in

Table 5.7: Ethno-Linguistic Fractionalisation and Growth

	(1) OLS	(2) step 2	(3) step 2
$\ln y_{t-1}$	-0.056		
	(0.013)		
ln(inv)	0.093		
	(0.014)		
ln(n+g+δ)	-0.183		
	(0.064)		
ln(schooling)	0.021		
	(0.013)		
Protestant		-0.063	
		(0.026)	
Eastern rel.		0.205	
		(0.071)	
ATiger		0.058	
		(0.027)	
ELF	-0.0008	-0.0003	-0.0004
	(0.0003)	(0.0003)	(0.0004)
m_1	3.47	4.39	4.73
m_2	1.35	3.16	3.58
adj R^2	0.21	0.07	0.01
F-test	p=0.00	p=0.00	

Note: Standard errors in parentheses.
n=483, 84 countries

Chapter 6. They use a measure of ethno-linguistic fractionalisation to proxy for ethnic diversity. The ethno-linguistic fractionalisation measure gives the probability that two randomly drawn people from one country do not speak the same language.[6] Easterly and Levine (1997) show that this measure of ethnic diversity is highly correlated with non-linguistic measures of social polarization such as the minorities at risk measures provided by Gurr (1993). Thus, they argue that ethnic diversity leads to rent-seeking behaviour which reduces the consensus for public goods, which in turn causes low growth rates.

[6] The original data was collected by Soviet researchers in the early 1960s. Mauro (1995) was the first economist to use this measure in a study of growth.

Here, we want to concentrate on one of the key Easterly and Levine (1997, 1998) results, namely that the coefficient on ethno-linguistic fractionalisation (ELF) is significant and negative in growth regressions. In Table 5.7 we add ELF to the augmented Solow model. Please note that for these regressions the sample size was reduced to 84 countries, because ELF is not available for Bangladesh. The OLS levels regression confirms the Easterly and Levine (1997, 1998) result, the coefficient on ELF is negative and significant at the one percent level. However, the coefficient on ethno-linguistic fractionalisation is not significant in our 'step two regression'. This suggests that high ethnic diversity is not associated with lower growth rates once we account for the presence of unobserved country specific effects in the estimation of the augmented Solow model.

5.5 Summary and Conclusion

In this chapter we have re-examined the relationship between growth and some measured country characteristics. We concentrate on truly time invariant characteristics such as geographical location, and cultural variables which only change very slowly over time, such as religious affiliations and ethnic diversity.

We show that the commonly found results in the literature were supported by our OLS levels estimation. If we neglect the presence of unobserved country specific effects and the endogeneity of investment in estimation, then the augmented Solow model appears not to fully account for low growth rates in Africa, Latin America, countries with a high proportion of Protestants and ethnically diverse countries; nor for the high growth rates experienced by the Asian Tiger economies and countries with a high proportion of people practicing Eastern Religions.

When we account for unobserved country specific effects and endogeneity using our 'two

step regression' procedure, we cannot confirm all of these findings. In the 'first step' of our estimation procedure, we use the preferred system GMM coefficient estimates of the augmented Solow model in order to calculate the residuals. In the 'step two regression' we investigate whether any of these measured country characteristics can account for some of the variation in these growth regression residuals. Our 'step two' regressions are estimated by OLS, and the estimated coefficients on the measured country characteristics are consistent if and only if all of these measured country characteristics are uncorrelated with the unobserved country specific effects. Since this is a strong restriction we only want to interpret these coefficient estimates as indicators of whether or not certain measured country characteristics are associated with high or low growth rates. We do not want to attach any causal significance to these results.

We find that the coefficients on the Africa and Latin America dummies as well as on the measure of ethnic diversity are insignificant. This suggests that the augmented Solow model can account for the low growth performance in Sub-Saharan Africa, Latin America and countries with high ethnic diversity if we allow for unobserved country specific effects and endogeneity in the estimation of the model's coefficients.

We do find that the augmented Solow model cannot account for the very high growth performance of the Asian Tiger economies. Furthermore, our results suggest that the proportion of Eastern Religions can account for some of the variation in the growth regression residuals, even when we include the Asian Tiger dummy.

We also find some evidence that countries with a high proportion of Protestants experienced lower growth rates than predicted by the augmented Solow model. As Barro's (1997) study suggests this may be due to institutional characteristics which these countries have in common. However, the investigation of the relationship between institutional characteristics and growth

is left for future research.

5.6 Appendix

5.6.1 Sample

The sample consists of :

22 Sub-Saharan African countries: Benin, Botswana, Cameroon, Central African Republic, Ghana, Kenya, Liberia, Malawi, Mali, Mauritius, Mozambique, Niger, Rwanda, Senegal, Sierra Leone, Sudan, Tanzania, Togo, Uganda, Zaire, Zambia, and Zimbabwe.

21 Latin American countries: Costa Rica, Dominican Republic, El Salvador, Guatemala, Haiti, Honduras, Jamaica, Mexico, Nicaragua, Panama, Trinidad and Tobago, Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Paraguay, Peru, Uruguay, and Venezuela.

3 Asian Tiger economies: South Korea, Hong Kong, and Singapore

39 other countries: Algeria, Australia, Austria, Bangladesh, Belgium, Burma (Myanmar), Canada, Denmark, Egypt, Finland, France, Germany, Greece, India, Ireland, Israel, Italy, Japan, Jordan, Malaysia, Nepal, Netherlands, New Zealand, Norway, Papua New Guinea, Pakistan, Philippines, Portugal, Sri Lanka, Syria, South Africa, Spain, Sweden, Switzerland, Thailand, Tunisia, Turkey, UK, USA.

Chapter 6

An Econometric Analysis of Civil War

6.1 Introduction

The literature on the analysis of civil war can be broadly categorized into two parts. On the one hand there is the theoretical analysis of the economics of conflict and on the other hand there is a large number of case studies mainly generated by political scientists, anthropologists and development studies scholars.

Several approaches have been suggested within the theoretical analysis of conflict. Hirshleifer (1987, 1994) suggests an analytical framework which is a combination of rational choice and game theory for the analysis of conflicts, which can for example be strikes, law suits, international or civil wars. In Hirshleifer's model the decision makers in the potential conflict situation are two individuals and, based on their preferences, opportunities and perceptions, they decide whether or not to start and continue a conflict. However, as Hirshleifer (1987, 1994) points out, strikes or wars are a collective activity, not only based on the decisions of two individuals. Thus, in a comprehensive analysis of civil conflict the processes which lead to

group action should also be modelled.

In Roemer's (1985) theory of revolutions two individuals compete for the support of the population. The revolutionary leader (Roemer calls him Lenin) attracts followers by promising general non-excludable benefits from a revolution against the incumbent ruler (the Tsar). The revolutionary leader must overcome the free-rider problem associated with non-excludable benefits from revolution. Roemer (1985) suggests that charisma of the revolutionary leader, together with revolutionary ideology and class identification are essential to overcome these free-rider problems.

Grossman (1995, 1999) develops a microeconomic framework of insurrections which is based on utility theory. The decision making agents in his model are the incumbent ruler, the potential leader of the insurgence and a large number of peasant or worker households. The incumbent ruler tries to recruit the households for soldiering activities in order to maintain the *status quo* and the insurgence leader tries to recruit households for his insurgence. Depending on their expected income, households decide whether to support the incumbent ruler or the leader of the insurgence. The expected income depends on returns to the different activities a household can engage in: home production, market production, soldiering and insurgence. In addition, the technology of insurrection and the discount rate will determine whether or not there will be an insurrection.

In contrast to these microeconomic studies, there is a large number of case studies of civil wars. Stewart, Humphrey and Lea (1997) provide an overview of the economic and social consequences of war, while Stewart (1998) gives a survey of the economic causes of civil conflicts. Bardhan (1997) provides a taxonomy of the catalysts of civil conflict and the study by Horowitz (1998) reviews recent political science and anthropological studies of ethnic conflicts.

Case studies have generated a lot of information about particular civil wars, however, Stewart (1998, p 2) as well as Horowitz (1998, p 13) conclude that no generalisations can be drawn from the case studies they reviewed. The political science study by Gurr (1993), on the other hand, uses data on minority groups in a cross-section of 127 countries and examines civil conflicts. His results suggest that political factors such as the level of democracy, state power and institutions can determine whether conflict takes the form of protest or rebellion.

In this chapter we propose a simple microeconomic framework which generates some hypotheses which can be tested empirically. For our empirical tests we use the comprehensive Singer and Small (1994) data set of wars, which provides global data on wars for 1816-1992. We draw on the analysis of Collier and Hoeffler (1998) and Collier, Hoeffler and Söderbom (1998).

Here we concentrate on three main questions. First, whether civil wars have economic causes. The subject has not, to our knowledge, previously been investigated econometrically. Second, we analyse the duration of civil war and third, the duration of postwar peace. We investigate whether civil wars are self-perpetuating and conversely, whether peace gradually becomes self-sustaining. Although this is of considerable importance for the effective timing of peace-promoting interventions, it has not to date been econometrically researched.[1] Nor is there a consensus of either practitioner or popular opinion. Do wars burn themselves out as the participants become exhausted, or become more entrenched as atrocities provoke sentiments of revenge? Does postwar peace becomes more secure as memories of war fade, or become increasingly fragile as the pressure cooker of old conflicts heats up again? These propositions clearly have opposing implications for the timing and duration of any peace-promoting intervention.

[1] A related study by Bennett and Stam (1996) investigates the duration of international wars in terms of political and military variables.

This chapter is structured in the following way. In section 6.2 we present a simple analytical framework proposing that the duration of conflict and that of postwar peace are determined by a common process which is considerably different to that which determines whether initially peaceful societies experience civil wars. In section 6.3 we discuss the variables used in the analysis, in particular the use of a comprehensive data set of civil wars (Singer and Small (1994)). Section 6.4 examines the causes of civil war and section 6.5 presents an analysis of the duration civil wars and postwar peace. Section 6.6 concludes by discussing the policy significance of the results.

6.2 Analytical Framework

Below we will show that the decision to rebel, as well as the decision to settle for peace, can be characterised by duration data methods. We focus on the hazard function, i.e. how the probabilities of a war ending (starting) depend on the elapsed time of war (peace) respectively, and on a set of explanatory variables. Using a proportional hazard model, we assume a functional form for these probabilities such that:

$$h = v(x)u(t) \tag{6.1}$$

where h denotes the hazard rate. Approximately the hazard rate h gives the probability of the war (peace) ending at time t conditional on the country being at war (peace) up to time t. Time t does not refer to calendar time, but denotes the length of the war (peace) spell. Thus, t and x is a set of regressors. We first consider the vector of regressors, and for this we draw upon a theory of the determinants of civil war. We then consider why the hazards of peace and war might differ from this model of the causes of conflict.

6.2.1 A Model of the Causes of Civil War

It obviously takes more than one agent to have a civil war, and the phenomenon is sometimes modeled as a two-agent game between a rebel group and a government (for example, Hirshleifer (1987, 1994 and Azam (1995)). In such an approach it is difficult to make the civil war outcome other than a negotiating mistake: both agents could do better by appropriate side-payments. However, in its origins civil war is not a two-agent game. No incipient non-government rebel movement can guarantee peace because it cannot prevent the emergence of other organisations. Nor need there be a common level of grievance which each incipient rebel group is representing, so that if the government satisfies one incipient group it satisfies all others. The aspirations of incipient rebels may be much greedier than the hope of equal treatment with other citizens, so that the government simply cannot satisfy the aspirations of all potential rebel groups. Prior to war, there may therefore be no rebel agent with which a government can negotiate a sustainable peace, so that the government may face the choice between war and abdication. If so, it will presumably choose war. Government leaders do not usually personally suffer from fighting, and in the event of defeat can usually escape at the last moment (Mengistu and Mobutu).

However, once a war is underway, one rebel organisation may acquire the key feature which turns civil war into a two-agent game, namely the ability to suppress rival organisations. Indeed, fighting between rebel groups is common in the early stages of civil war as one group suppresses its rivals (e.g. among Eritrean rebels against Ethiopia and among Sudanese rebels against the Sudanese government).

Starting from peace, the cause of civil war is not the breakdown of a two-agent game. Government actions may reduce or increase the risk of war, but war need not be a negotiating error. In some societies the representative incipient rebel movement may simply have a very

strong incentive to resort to violence. However, civil war only occurs if incipient rebel groups are able to organise themselves on a large scale. Political terrorism can be undertaken by a handful of people, or even an individual, but civil war can only be produced by mass-collaboration in political violence. In determining why some societies have civil wars and others do not, we must therefore focus upon the determinants of whether incipient rebellions escalate. That is, upon the decisions of young men to enlist in incipient rebel movements. It is these decisions which determine whether such movements never develop beyond an irritant, or grow into organisations which can challenge the government for control of territory.

Consider the decision to enlist in a rebel movement. The decision has a dual feature: it is both a political act and an economic act. As a political act it is analogous to a vote: the enlisting agent registers support for a political organisation. People register their support partly to influence political choices, because they have either national or distributional policy concerns. They also register support as a pure expression of group allegiance, analogous to supporting a football team. However, as an economic act the enlistment decision is analogous to enlisting in a conventional army. The agent is joining a rebel fighting force as an employee. Further, the employment decision is quasi-irreversible: whereas joining a rebel army is (usually) a voluntary act, choosing to withdraw from it is classified as desertion and is severely punished. In the limit, as with many recruits to a government army, the cause for which the rebellion is ostensibly being fought may be entirely peripheral to the true motive for enlisting.

Thus, we model the causes of civil war as a function of the benefits and costs of rebellion. The benefits of rebellion are contingent upon rebel victory, the probability of which, p, decreases with the military capability of the government. We proxy the military capability of the government relative to the rebel population by the per capita taxable base of the economy as measured by

GDP per capita (Y). Contingent upon victory, the benefit to the rebels may be the capture of the state. Thus, the benefit is a function of the taxable base.

We propose two costs of rebellion, its opportunity cost and the cost of coordination. The opportunity cost is the loss of income sustained during the conflict. We proxy this by per capita GDP. Coordination costs are the transactions costs associated with collective action. The difficulty of coordinating rebel activity is proxied by ethnic composition. We postulate that in ethnically homogenous societies rebels face difficulties of differentiating themselves from government supporters, whereas in highly fractionalised societies they face difficulties agreeing with each other. The ideal ethnic composition for rebellion is one in which there are two ethnic groups, only one of which is identified with the government. Ethnic composition is measured by the index of ethno-linguistic fractionalisation (ELF) which scales composition on the range 0-100. Since coordination costs are postulated to be at their minimum in the middle of this range, the index is entered as a quadratic.

The expected utility of rebellion depends on the duration of the rebellion, thus

$$E(U_W) = \int_0^\infty (U_W|D) \cdot g(D) \cdot dD$$

where $g(D)$ denotes the probability density function of the rebellion's duration. In more detail we want to specify the expected utility of rebellion as

$$E(U_W) = \int_0^\infty \left[\int_D^\infty \frac{p(Y|D) \cdot G(Y)}{(1+r)^t} dt - \int_{t=0}^{t=D} \frac{(f(Y)+C)}{(1+r)^t} dt \right] \cdot g(D) \cdot dD \qquad (6.2)$$

where

p = probability of rebel victory, and p'<0

G = gain conditional upon victory

D = expected duration of warfare

Y = per capita income

C = coordination costs

r = discount rate

U_W= rebel utility

The decision to rebel will be taken if $U_w > 0$. The probability of rebellion in a country will be a function of Y, C and the expected duration, D. A theory of the probability of civil war is directly a theory of the duration of peace, since a low probability of war per period will translate into a longer predicted duration of peace. Indirectly, it also provides the basis for a theory of the duration of conflict. The maximum duration of rebellion at which, if it is correctly anticipated, rebellion remains rational, D_{max}, is a function of the other two variables:

$$D_{\max} = f(Y, C) \tag{6.3}$$

If the duration is correctly anticipated and only rational rebellions occur, then the actual duration of a rebellion will be bounded over the range $\{0, Dmax\}$. The mathematical expectation of D, $E(D)$, will be the integral of the density function $g(D)$ multiplied by D over this range and so will be an increasing function of D_{max}. Hence,

$$E(D) = g(Y, C) \tag{6.4}$$

Equation (6.4) constitutes a baseline prediction for the determinants of the duration of civil war and postwar peace, and we first state it as a testable hypothesis:

Hypothesis 1: the duration of civil war will be a function of the same variables as the causes of civil war, namely Y and C.

However, we now introduce three factors which predict that hypothesis 1 will be rejected. They are: errors in expectations, irreversibility in decisions, and changes in the costs of conflict.

6.2.2 Errors in Expectations

As with other areas of expectations, a reasonable starting assumption is that potential rebels have rational expectations about relative military capabilities. However, given this assumption, and the assumption that the expectation of the military outcome influences whether rebellions occur, it follows that on average in actual rebellions, rebels will have been over-optimistic. Potential rebels with pessimistic expectations will be less likely to fight than those with perfect foresight, whereas potential rebels with optimistic expectations will be more likely to fight. Hence, actual rebellions will tend to be motivated on average by systematic over-optimism: wars will turn out to last longer than rebels initially expected. This is a similar problem to the winner's curse as discussed in the auctions' literature.[2]

These errors are corrected once the rebellion starts and the rebels learn about relative military capabilities. On average, rebels learn that they have been over-optimistic. As a result, some rebellions will end quickly, either because the rebels have been so over-optimistic that they are quickly defeated, or because a negotiated settlement is found to be much more attractive once it is realised that rapid victory is unattainable. There is no equivalent effect during the

[2]The best know example of the winner's curse is bidding for offshore oil tracts. Even if the competitors' bids are unbiased estimates of the value of the oil tract, the winner of the auction is likely to have overestimated the value of the oil tract. Thus, the winner is likely to lose money on the investment if the competitors do not downscale their bids. This is known as the winner's curse. For further details see for example Milgrom and Weber (1982) and the literature quoted there.

early years of a postwar peace. Typically, in this phase both parties to the conflict retain the possibility to reactivate their prior military capabilities, and so retain knowledge of their relative standing. The correction of errors at the start of wars arises only if they follow long periods of peace.

Hence:

Hypothesis 2: The hazard of peace during a civil war will initially be high and then fall.

Hypothesis 3: There will be no analogous abrupt change in the hazard of renewed war following a peace.

6.2.3 Irreversible Decisions

We now consider more closely the 'production' of rebellion. A successful rebellion requires that a fighting force should first be recruited and then sustained over the required period of military engagement.

Initial recruitment depends upon the decisions of those who might potentially join a rebel army. The most important factor in the decision whether to join is likely to be the personal opportunity cost of joining, as proxied by per capita income. Hence, a potential rebel leader faces the opportunity cost of the time of potential recruits as an obstacle to the initial formation of an army.

Conditional upon successful initial recruitment of forces, the rebel leader now faces a different problem. Potentially, the leader may face the recruitment problem as a continuous constraint, since recruits may reassess their opportunity cost and so leave the army. This is obviously at its most severe at times of intensified fighting. Ideally, a recruit would like to get the benefits of group membership without the risks of battle. All armies face this problem and

all adopt the same solution, which is to restrict the choice of exit. The ability to wield force constituted by an army is used, in part, upon its own members. Exit becomes classified as desertion and is severely punished. Similarly, the force constituted by the army can be used to coerce subsequent recruitment, a phenomenon common in many civil wars. An implication of this is that whereas the opportunity cost of time is an important constraint upon the initial formation of a rebel army, it is much less of a constraint upon the capacity for sustained fighting conditional upon formation.

Hypothesis 4: per capita income will be less important in determining the duration of war than in determining its cause.

However, in order to use the army coercively against its own members, it has to be organised into a decentralised hierarchy: privates must be subject to the discipline of officers who are sufficiently proximate to supervise their conduct and detect attempted escape. This command structure, which is again common to all armies, carries a new danger. The army in aggregate is, in effect, a coalition of its local commanders, since the troops in each locality take their orders from the commander in each locality. Armies defend themselves from mass individual desertion at the price of group defection. Facing a rebellion, a government will therefore attempt to induce rebel officers either to defect with their forces, or splinter the rebellion by forming a faction opposed to the initial leader. This is, indeed, how many rebellions fall apart, examples being the White Army in Russia, during 1920, the southern Sudanese during the 1990s, and the Kurdish opposition to Iraq in 1996. The use of a coercive hierarchy thus transforms the nature of the problem faced by the rebel leader from the constraint of individual exit, determined by the opportunity cost of time, to that of the maintenance of cohesion. The capacity to sustain rebellion, and hence its maximum feasible duration, becomes a function of the costs of rebel

coordination, as proxied by ethno-linguistic diversity.[3] Hence:

Hypothesis 5: coordination costs will be more important in determining the duration of war than in determining its cause.

As a consequence, the occurrence of rebellion and the duration conditional on occurrence, will not be determined by the same process. In the limit, the coordination problem may be unimportant in the initial recruitment process because those potential rebels with the lowest opportunity cost join even without knowing the decisions of others, whereas those with higher opportunity costs join in the knowledge of the decisions of those with lower costs. Sequencing of recruitment by opportunity cost can thus overcome the absence of decision coordination. Hence, the occurrence of rebellion would be influenced by costs of time but not by costs of coordination. Conversely, the capacity to sustain rebellion might, through the use of the coercive power of the rebel army, entirely eliminate the importance of the opportunity cost of time, but replace it with the problem of officer coordination. In practice, such a complete change in causal factors is unlikely. The formation of a rebel army will be hindered by a high degree of fractionalisation. Conversely, exit from the army, even though subject to severe punishment, will still occur and be influenced by the opportunity cost of time. For example, in the largest civil war of the twentieth century, that of Russia 1919-20, there were four million deserters despite the death penalty, and the rate of desertion was up to ten times higher in summer than in winter because of the high opportunity cost of peasants being absent from their farms at harvest time. Thus, both the occurrence and the duration of civil war can be expected to be functions of the same variables, but the relative importance of the variables is predicted to be different.

[3] An indication of this is that successful military leaders in ethnically diverse societies commonly overcome the problem of coordination in the officer core by confining it to a single ethnic group, although this has obvious costs in terms of army morale and recruitment, examples being Iraq, Zaire and Frelimo.

Now consider the hazard of renewed war in the early years following settlement of a conflict. The parties to a civil war usually retain their military capability during the early years of a postwar peace. As a result, the risk of renewed war, and hence the duration of the postwar peace, is more analogous to the maintenance of the previous war than to the maintenance of peace in a society which has not experienced civil war for a long time (so that a rebel army must be recruited from scratch). Hence:

Hypothesis 6: the maintenance of postwar peace will be determined by the same process which maintained the previous war, rather than the process which sustains peace in societies which have not recently experienced civil war.

6.2.4 Changing Costs of Conflict

We now consider why the costs of conflict might change as the duration of a civil war increases.

The first time-dependent effect is sunk costs. From (6.2) all the costs of war accrue before any of the benefits. The sunk cost effect diminishes the probability that the decision to rebel will be reversed as time adds to the information stock, so that the hazard of peace will diminish with the duration of the war.

A second possible effect which produces the same result is the gradual polarisation of the population. This may occur partly because many people are forced to take sides, and partly because atrocities create a widespread desire for revenge. Keely (1996) shows that in stateless societies the main motivation for warfare is the desire for revenge for previous acts of warfare. The effect of polarisation is to reduce coordination costs for the rebels as the war proceeds: a part of the population comes to identify itself as hostile to the government, and as joined in

common purpose behind a rebel organisation and leadership.[4] Formally,

$$C = C(D), C_D < 0 \tag{6.5}$$

As with sunk costs, polarisation cumulatively reduces the cost of continued fighting as the war proceeds and so continuously reduces the chance of decision reversal. It therefore produces a downward-sloping hazard function.

A final possible effect of duration is the exhaustion of the combatants. For example, at the level of the ordinary combatant the very long war in Mozambique is said to have left participants with a desire for peace. The negotiated end to the conflict had provision for a substantial army of national unity made up of the combatants from each side. In the event the post-war army had to be drastically scaled down because a large majority of the combatants opted to leave their armies and go back to peacetime activities. At the level of the leadership, it has been suggested that one motivation for Yasser Arafat to conclude a peace settlement with Israel was that he wished to be head of a Palestinian entity before he died. While sunk costs and polarisation produce a downward sloping hazard function, exhaustion produces a rising hazard function. Hence:

Hypothesis 7: the net effect of duration on the hazard of peace is a priori ambiguous.

The hazard of renewed civil war following a peace settlement is potentially subject to change as a result of both the polarisation and exhaustion effects fading. The former would produce a declining hazard of war and the latter a rising hazard. Hence:

Hypothesis 8: the net effect of duration on the hazard of renewed war is a priori ambiguous.

[4] One way of thinking of this is that the society gradually accumulates a stock of social capital appropriate for civil war.

Cumulatively, these eight hypotheses form a set of testable propositions to which we now turn.

6.3 Data

In the subsequent analysis, the dependent variables are the duration of civil war and the duration of postwar peace. We use and update the Singer and Small (1994) data set on civil wars from 1816-1992 which dates the start and end of civil wars by month.[5] Hence, our durational data is in months. Singer and Small (1982) provide an operational definition of civil war. The authors define wars in terms of violence, not in terms of the goals of the protagonists or the results of the war. A civil war in Singer's and Small's typology is based on four dimensions. First, one of the primary actors in any conflict identified as a civil war must be the national government in power at the time hostilities begin. Secondly, the concept of war requires that both sides have the ability to inflict death upon each other. As a rule of thumb Singer and Small (1982) define that in a civil war the stronger forces must sustain at least five per cent of the number of fatalities suffered by the weaker forces. Thirdly, significant military action must take place. Only civil wars that resulted in at least 1000 battle related deaths per year are included in the data set. This figure includes civilian as well as military deaths. Fourthly, the war must be internal to the country. On the Singer and Small definition of internality this produces some important exclusions. Wars which they regard as being between a country and dependent territories, such as those in Angola, Mozambique and Eritrea prior to formal independence, are classified by Singer and Small not as civil wars but as a sub-category of international wars termed 'extra-systemic'. However, since they are in many respects more akin to civil wars,

[5] Please refer to the appendix for details on how we updated the Singer and Small (1994) data set.

being fought entirely within national boundaries, for our purposes we have included them in our sample.

The Singer and Small definition of civil war is not scale-neutral, because it imposes an absolute threshold of one thousand deaths per annum as a minimum. Since the size of countries differs so enormously, it is not realistic to rescale the Singer and Small definition so that the threshold would be proportionate to the population. For example, if the threshold for civil war in the Turks and Caicos was set at two deaths during the year, which is obviously ridiculously low, China would not meet the requirement even were civil conflict to kill a quarter of a million people. Hence, the absolute threshold utilised by Singer and Small has much sense to it, even though it makes large countries more prone to civil war. There are other reasons why population size might affect the probability of civil war[6], but since these cannot be disentangled from the statistical threshold effect, there seems little value in considering them. To allow for the threshold effect, population size will be entered as an explanatory variable in all regressions.

We proxy the transactions costs of coordinated action by cultural distinctness. Cultural distinctness is measured by an index of ethno-linguistic fractionalisation. This variable measures the probability that any two citizens will be drawn from a different ethno-linguistic group. The variable is re-scaled so that complete homogeneity scores zero and maximum fragmentation scores 100. It was first utilised by Mauro (1995) and is measured for most countries in the early 1960s.

[6] One argument why population size may matter for rebellion is that benefit of rebellion may either be the capture of the state or secession from it. The larger the population size, the stronger may be the wish to succeed from the state. (Please refer to Alesina and Spolaore (1997) who present a model of the optimum size of a nation.) On the other hand the costs of coordination of a rebellion depend positively on the size of the population. A rebellion covering a given proportion of the population (and thus, *ceteris paribus*, standing the same change of success), will require communication between a larger number of people in a country with a larger population. Thus, a greater population may be associated with a lower probability of coordinating a rebellion.

We use World Bank data for the population variable and PPP adjusted per capita GDP from the Penn World Table as the income variable.

While the series built by Singer and Small gives the potential for an analysis over a period of more than a century, data on per capita income and the ethno-linguistic index is only available for more recent periods. Hence, in the war models we concentrate on civil wars which started after 1960. With this restriction the hazard of peace while at war can be estimated on a sample of 45 spells. Using only data from 1960-96 severely reduces the sample size for the hazard of renewed war following a peace settlement. Thus, we use data from 1945-96 for our peace hazard model which increases the sample size to 56 spells.

6.4 The Causes of Civil Wars

From equation (6.2), the occurrence of civil war is posited to be a function of the economic variables Y and C, proxied by per capita income and the index of ethno-linguistic fractionalisation (and its square). As discussed above we also control for the threshold effect by adding population size as an explanatory variable. We first test these propositions in a probit model.

The risks of a civil war continuing are additionally posited to be functions of the prior duration of the war. Similarly, we can test for the continuation of postwar peace as a function of the prior duration of peace and economic variables. These durational effects can only be tested using a hazard function. For the 'hazard' of peace as a war continues the relevant sample is of countries which are initially at war. For the hazard of renewed war in a country which has reached a peace settlement the relevant sample is of countries which are initially in postwar peace.

By contrast, the probit extends the data set of the hazard function by adding all those

Table 6.1: Economic Variables and the Occurrence of Civil Wars 1960-96

Variable	Coefficient	S.E.	z-value
constant	-0.777	0.528	-1.47
GDP_{60}	-0.56E-3	0.20E-3	-2.83
ELF	0.38E-1	0.22E-1	1.73
ELF^2	-0.43E-3	0.25E-3	-1.75
$lnPop_{60}$	0.385	0.129	2.99

		predicted	
		0	1
actual	0	66	6
	1	16	12

Note: N=100, Pseudo R^2=0.26

observations in which the duration of war is zero. The probit can thus be estimated on the much larger sample of all countries which are initially at peace (100), whether or not they subsequently experience war. In the probit the dependent variable takes a value of zero if the country experienced no civil war during 1960-96. It takes a value of one if the country experienced civil war during the period, regardless of the length of the war and whether the country experienced single or multiple spells of civil war. 28 countries experienced at least one civil war during 1960-96.

In Table 6.1 we report the results of the probit regression for the occurrence of war. The sample period is 1960-96. Per capita income, ethno-linguistic fractionalisation and population size are measured at the beginning of the period. They all have significant effects on the risk of civil war. Potentially, per capita income could either reduce or increase the risk of war since it proxies not only the opportunity cost of rebellion and the capacity of the government to defend itself (both of which reduce the risk), but also the potential gain to capturing the state (which increases the risk). However, the net effect is dominated by the former of these: countries with

higher per capita income had a lower risk of a war occurring. As expected, the coefficient on population size is positive. Countries with larger populations were more prone to civil war, as suggested by the absolute threshold definition of civil wars. The effect of ethnic diversity on the risk of civil war is non-monotonic. We interpret the regression results as evidence that rebel coordination would be more difficult both in societies in which the entire population was from the same group, so that there was no obvious distinction between government and rebel supporters, and in societies which were so highly fractionalised that rebellion required coordination across multiple distinct groups. Coordination costs are at their lowest when the population is polarised between an ethnic group identified with the government and a second, similarly sized ethnic group, identified with the rebels. The risk of war is at a maximum when the ethno-linguistic index takes a value of about 47.5.

6.5 The Duration of Civil Wars and Postwar Peace

We now turn to the investigation of durational effects through the hazard function. As indicated in section 6.2 we specify the hazard as a product of two terms, $u(t)$ and $v(x)$. A common assumption in duration studies is that $u(t)$ follows a Weibull distribution, which implies that the hazard rate can be specified as:

$$h_i(t, X, b, p) = p \cdot [\exp(-X_i^{'} b)]^p \cdot t^{p-1} \tag{6.6}$$

where b is the vector of coefficients associated with the regressors, and p is the duration dependence parameter. There will be positive (negative) duration dependence for p larger (smaller) than unity, and no duration dependence if p equals one. The Weibull model has recently been used to characterise the determinants and the duration dependence of interstate wars (Bennett

and Stam (1996) and Vuchinich and Teachman (1993)).

The Weibull specification is analytically tractable, but also quite restrictive. First, unobserved heterogeneity is omitted, which may cause spurious negative duration dependence (see e.g. Lancaster (1990, chapter 4)). Although (6.6) can be extended to control for heterogeneity, the methods for doing so are not uncontroversial. Second, the baseline hazard in (6.6) is restrictive in that it imposes a monotonic duration dependence, which may lead to seriously misleading inference: if the actual hazard function is u-shaped, for instance, Weibull duration dependence estimates will be upward biased for short durations and downward biased for long durations. Similarly, the Weibull specification is unsuited to testing the step-function hazard predicted by hypothesis 2.

A more flexible stance to characterising duration dependence can be taken by adopting a semi-parametric estimator. Cox (1972) developed a semi-parametric model within the proportional hazards framework, which was modified by Prentice and Gloeckler (1978) to allow for stepwise estimation of the baseline hazard. The Prentice-Gloeckler model can be written as

$$h_i(t, X, b, h_0) = \exp(-X_i^{'} b) h_0(t) \tag{6.7}$$

where h_0s are the stepwise baseline hazard probabilities to be estimated. Thus, in contrast to the Weibull model, the shape of baseline hazard is totally flexible. In principle it is straightforward to extend the model to control for unobserved heterogeneity (see Meyer (1990)), but identification can be problematic in small samples. This was the case for our analysis. However, there is some evidence in the applied literature that controlling for unobserved heterogeneity has only minor effects on the hazard estimates as long as the baseline hazard is sufficiently flexible (e.g. Han and Hausman (1990); Bover, Arellano and Bentolila (1997)). Indeed, we

are not aware of any reference in which the baseline hazard is flexible and where unobserved heterogeneity had important effects on either the baseline hazard estimates or the estimates of the coefficients.

We first consider the hazard of peace during a civil war, the results being reported in Table 6.2.[7] Income and population are measured before the start of the civil war. Ethno-linguistic fractionalization is only measured around 1960, however, it is a slow changing variable and therefore applicable to the entire period. The Weibull model is specified in continuous time, and we use monthly data in the estimation.[8] None of the regressors is significantly different from zero, but qualitatively we obtain the inverse u-shape effect on ethno-linguistic fractionalisation predicted in section 2. This suggests that either a low or a high value of ethno-linguistic fractionalisation increases the chance of peace. The duration dependence parameter p is significantly less than unity, implying that the hazard is downward sloping.

To estimate the semi-parametric hazard model it is necessary to divide time into a number of intervals. Due to the small sample size it is desirable to keep the number of parameters to be estimated at a minimum, at the same time as allowing for a sufficient number of hazard probabilities. We have chosen annual interval, and we censor any spells longer than ten years at ten.

The results differ significantly from the causes of civil war as found in the probit. First, neither income nor population affect the hazard of the war ending, although both were significant in causing war. Secondly, the effect of ethno-linguistic fractionalisation is similar to the probit

[7] A positive sign on a coefficient associated with an explanatory variable implies that the hazard is shifted downwards for an increase in the variable, i.e. that the hazard probablities fall and that the expected duration increases.

[8] Singer and Small provide daily data on the start and end of wars. For our purposes it is sufficient to use monthly accuracy in the continuous time model.

of the causes of war. Thus, consistent with hypotheses 4 and 5, income is more important relative to coordination costs in causing war, but less important in determining its duration. However, the point estimates of the ethno-linguistic fractionalisation coefficients, the probability of the war ending is at its lowest for an index of 43. Hazard probabilities are monotonic transformations of the coefficients on d1-d10, and we plot these probabilities, along with the Weibull hazard, in Figure 6.1.[9] As predicted by hypothesis 2, the semiparametric hazard falls sharply after the first year of civil war. Thereafter, it remains relatively flat, implying that the effect of exhaustion offsets those of sunk costs and polarisation.[10]

We now turn to the hazard of renewed war in a society which has just achieved peace. The results are reported in Table 6.3. Since the number of peace spells was small for 1960-96, we added data from 1945-60 to estimate the peace hazard. This expansion of the data set is slightly problematic, because we have only have data for the right hand side variables from 1960 onwards. We use the values for the explanatory variables which are closest to the beginning of the peace spell.

The hazard models of peace ending are a hybrid between the causes of war after long periods of peace and the determinants of the duration of war.

As before the Weibull models is specified in continuous time and we use monthly data in the estimation. Ethno-linguistic fractionalisation and its square are marginally significant, confirming the inverse u-shape effect. Homogenous as well as highly fractionalized societies have a low risk of civil war, while societies which consist of two similarly sized groups have the

[9] The hazard estimates are obtained by the transformation $h(s) = 1 - \exp(\exp(x'b + d(s)))$. Hence, d1-d10 represent h1-h10 monotonically. For Figures 6.1 and 6.2 we use mean values of the regressors.

[10] Using a Wald Test, the hypothesis that the hazard rate in the second year is the same as that in the first year of the war was rejected at 10% level of significance. Individual tests of whether any of the coefficients associated with d3-d5 is equal to the coefficient on d1 were all rejected at 5% level, as was the joint hypothesis that the hazard rate during years two to five of the civil war is the same as during the first year of war.

Table 6.2: Hazard of War ending

	Semiparametric			Weibull		
	Coeff.	S.E.	z-value	Coeff.	S.E.	z-value
d1	-0.442	0.624	-0.709			
d2	-1.642	0.839	-1.957			
d3	-1.949	0.916	-2.126			
d4	-2.584	1.156	-2.235			
d5	-2.556	1.163	-2.199			
d6	-1.767	0.907	-1.948			
d7	-2.201	1.167	-1.887			
d8	-1.400	0.921	-1.520			
d9	-0.487	0.796	-0.621			
d10	-1.689	1.172	-1.441			
const				3.756	0.898	4.184
ELF	0.053	0.030	1.775	0.055	0.040	1.368
ELF^2	-0.62E-3	0.33E-3	-1.850	-0.001	0.000	-1.446
lnPop	0.146	0.184	-0.795	0.181	0.243	0.745
GDP	-0.16E-3	0.19E-3	-0.852	-0.2E-3	0.3E-3	-0.843
p				0.693	0.096	
Log-L		-80.649			-82.370	
N		45			45	

Note: Intervals are one year. Log likelihood values of the Weibull and the semiparametric models are not comparable, since the former model uses data in continuous time and the latter uses discrete time data.

Table 6.3: Hazard of Peace Ending

	Semiparametric			Weibull		
	Coeff.	S.E.	z-value	Coeff.	S.E.	z-value
d1	-3.727	0.838	-4.446			
d2	-3.045	0.765	-3.981			
d3	-4.819	1.200	-4.016			
d4	$-\infty$	–	–			
d5	-4.669	1.197	-3.900			
d6	-3.457	0.872	-3.964			
d7	-3.768	0.958	-3.935			
d8	-3.544	0.944	-3.753			
d9	-4.134	1.175	-3.519			
d10	-3.291	0.929	-3.542			
d11	-3.667	1.138	-3.222			
d12	$-\infty$	–	–			
d13	$-\infty$	–	–			
d14	-3.607	1.143	-3.155			
d15	-3.412	1.151	-2.964			
const				7.456	0.759	9.825
ELF	-0.048	0.031	-1.553	-0.053	0.034	-1.559
ELF^2	0.001	0.000	1.639	0.001	0.000	1.680
lnPop	-0.202	0.190	-1.060	-0.366	0.201	-1.818
GDP	0.000	0.000	-0.069	0.000	0.000	0.250
p				0.903	0.145	
Log-L		-88.4374			-67.768	
N		56			56	

Note: Intervals are two years. Log likelihood values of the Weibull and the semiparametric models are not comparable, since the former model uses data in continuous time and the latter uses discrete time data.

highest risk. As in the probit, but unlike the peace hazard, a greater population increases the risk of civil war. However, income remains insignificant. The duration dependence parameter p is not significantly different from unity, indicating that the probability of a war breaking out does not depend on the duration of peace.

For the semiparametric hazard model we divide time into biannual intervals. The inverse u-shape effect of ethnic diversity on the probability of civil war is confirmed. The probability of civil war is again at a maximum when the ethnic diversity index takes a value of 43. Neither income nor population are significant. These results which are only based on countries which have recently experienced a war can be contrasted with those from the probit which is dominated by countries which have been at peace for a long time. Hence, consistent with hypothesis 6 income is important in the process of maintaining peace in societies which have not recently experienced a war whereas in post conflict situations only ethnic diversity is important. We present the hazard probabilities along with the Weibull hazard in Figure 6.2. Neither the semiparametric hazard nor the Weibull hazard follow a trend, indicating that the probability of civil war is duration independent.[11]

This result may be due to the small number of observations and we decided to use a larger sample to analyse the duration dependency of peace spells. In this larger sample we use all the information provided by the Singer and Small data set and not only for the period after World War II. We include all countries which experienced at least one war between 1816 and 1996. As before postwar peace spells are measured as the number of months after a civil war. Since some countries experienced a number of civil wars during the period we list several peace

[11] The only two hazard rates of the semiparametric model that are in the neighbourhood of being significantly different from each other are those in the second and third periods. Using a Wald test, the hypothesis that they are the same is rejected at 10%, but not at 5%.

spells for these countries and our data set consists of 256 spells. Since we stop recording the duration of civil wars in 1996 we treat the data as single spell data with censoring. We do not have sufficient data on neither economic nor ethnic characteristics for years before 1960, thus, we estimate a Kaplan-Meier hazard function with this data. As before we group the peace duration data in biannual intervals. The hazard rate h(t) gives the number of countries in which a war broke out during time period t divided by the number of countries at peace during time period t. The hazard function for postwar peace using data from 256 spells is pictured in Figure 6.3. Within the first six years of postwar peace, the hazard of experiencing a new civil war is halved from about 0.12 to 0.06 and remains relatively constant thereafter (estimates for later years are relatively noisy due to fewer observations in later years). Note that 0.06 was also the hazard rate given by the (flat) hazard function using data from only post World War II spells. However, according to the Kaplan-Meier estimates the hazard of peace ending is duration dependent, the longer the peace lasts, the less likely is a renewed outbreak of civil war.

In Figure 6.4 we combine the results obtained from the war and peace duration models. This figure illustrates the hazard of peace and renewed war for a country which experienced a civil war which lasted for five years. Thus, during the first five years this hypothetical country is at war and at peace for all subsequent years. Combining our peace hazard results and war hazard results we can describe the stylized hazard of peace and renewed war of such a five year war as follows. According to our results obtained from the proportional hazard of peace model, we suggest that the risk of continuing the civil war is lower in the first year, but then rises and remains constant. After the peace settlement in the fifth year of the civil war, our Weibull results from the hazard of war model suggest that the risk of a renewed civil war is duration

independent.

6.6 Conclusion

As a starting point we set out a simple utility based model of civil war. Rebels will conduct a civil war if the perceived benefits outweigh the costs of rebellion. Using a probit model we test these propositions empirically. Furthermore, we suggest that the same explanatory variables may also be related to the duration of civil war. We test this suggestion in a proportional hazard model. Our results suggest that both the causes and the duration of civil war are related to a few socio-economic variables, however, the two processes are distinct. Income levels and population size are important causes of civil war, but have no effect on the duration of conflict and little effect on the duration of a post-conflict peace. By contrast, ethno-linguistic fractionalisation is both absolutely and relatively more important as a determinant of the duration of war and of postwar peace than as a cause of war, although in all cases its effect is non-monotonic: both homogenous and highly fractionalised societies are substantially safer than those with middle range diversity. We explained this non-monotonic relationship in terms of the costs of coordination of rebellion, which are minimised when society is split between two ethnic groups. We explained the differences between the causes and duration of civil war in terms of the irreversibility of the military recruitment decision and the problems posed by maintaining cohesion of the officer core during a conflict. The hazard of peace during a war falls sharply after the first year. We explained this in terms of the systematic over-optimism of actual rebels which is analogous to the winner's curse in auction theory.

Our findings suggest that raising incomes may only be an effective strategy for reducing the risk of civil war in societies which have not experienced one for a long time, but is likely to

be ineffective both during and after a conflict. Once the conflict has persisted beyond the first year the chances of resolving it sharply diminish and remain fairly constant. Hence, income enhancement efforts in the early years of peace may not be effective either in reducing the risks during these years, or in getting the society over an imagined high risk phase. In multi-ethnic conflict ridden societies in which single ethnic political entities are not feasible, changes which reduce the extent of ethnic diversity would tend to exacerbate the risks of renewed or continued civil war. A high level of fractionalisation seems more effective in shortening war and sustaining peace than income growth.

6.7 Figures

Figure 6.1: Hazard of War Ending

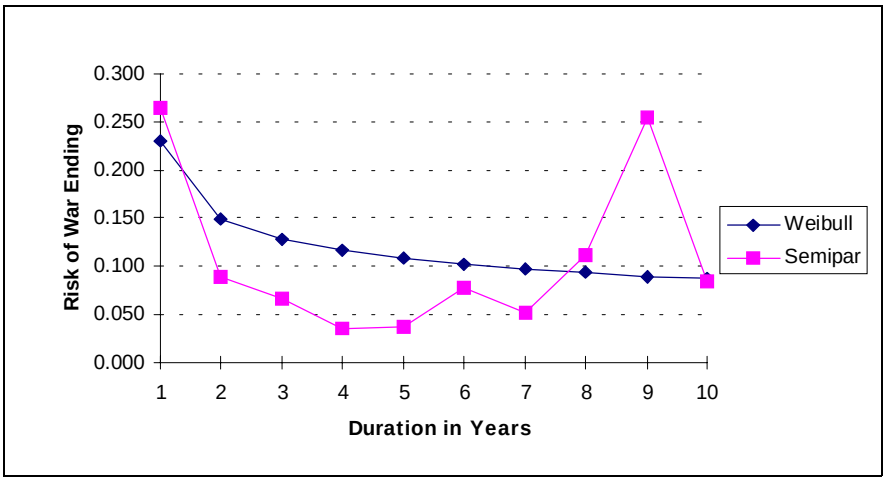


Figure 6.2: Hazard of Peace Ending

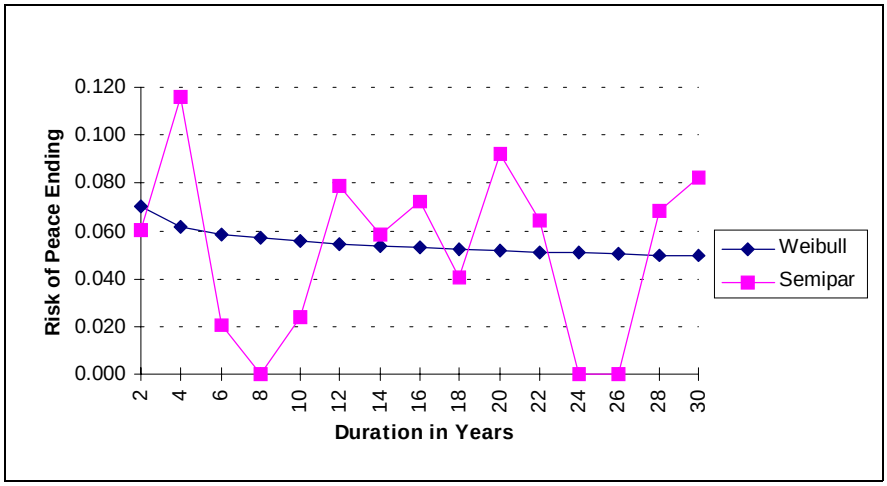


Figure 6.3: Hazard of Peace Ending - Kaplan Meier Estimates

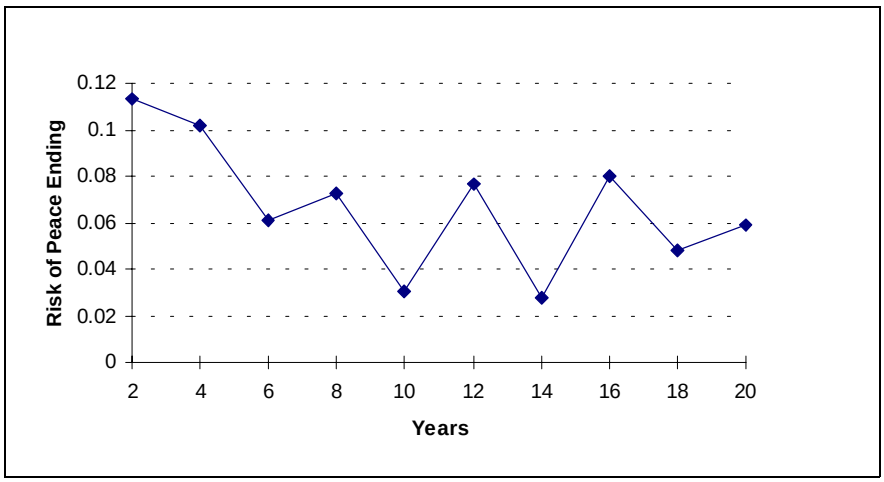
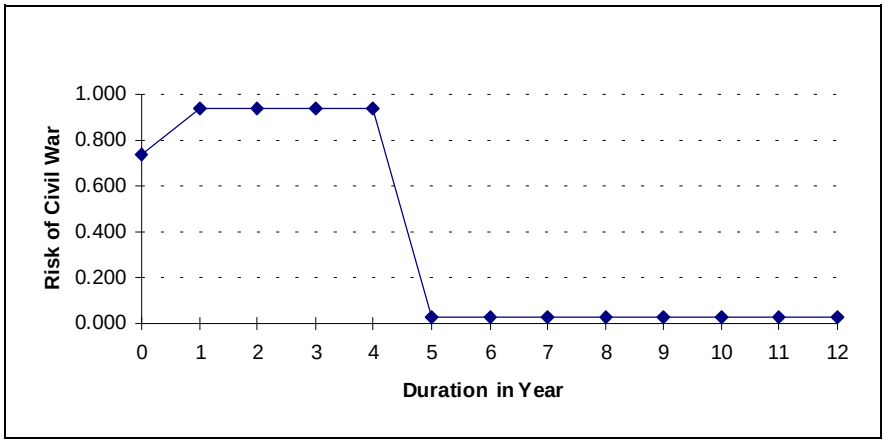


Figure 6.4: Duration of Civil War and the War Hazard after a Peace Settlemen



6.8 Appendix

6.8.1 Sample

Sample for the probit includes the following 100 countries:

Algeria*, Argentina, Australia, Austria, Barbados, Benin, Bolivia, Brazil, Burkina Faso, Burma* (Myanmar), Burundi*, Cameroon, Canada, Central African Republic, Chad*, Chile, Colombia,* Congo, Costa Rica, Denmark, Dominican Republic*, Ecuador, Egypt, El Salvador*, Ethiopia*, Finland, France, Gabon, Gambia, Germany, Ghana, Greece, Guatemala*, Guyana, Haiti, Honduras, Hong Kong, Iceland, India*, Indonesia*, Iraq*, Ireland, Israel, Italy, Ivory Coast, Jamaica, Japan, Kenya, Korea, Liberia*, Madagascar, Malawi, Malaysia, Mali, Malta, Mauritania, Mauritius, Mexico, Morocco*, Mozambique*, Nepal, Netherlands, New Zealand, Nicaragua*, Niger, Nigeria*, Norway, Pakistan*, Panama, Papua New Guinea, Paraguay, Peru*, Philippines*, Portugal, Saudi Arabia, Senegal, Sierra Leone*, Singapore, Somalia*, South Africa, Spain, Sri Lanka*, Sudan*, Sweden, Switzerland, Syria, Tanzania, Thailand, Togo, Trinidad and Tobago, Tunisia, Turkey*, UK, USA, Uganda*, Uruguay, Venezuela, Zaire*, Zambia, Zimbabwe*.

An asterisk indicates that the country experienced a civil war during 1960-92.

The hazard of a war ending used data from the following 45 civil wars

Algeria (2), Angola (1), Burundi (3), Chad (1), Columbia (1), Dominican Republic (1), El Salvador (1), Ethiopia (1), Guatemala (2), India (1), Indonesia (1), Iran (2), Iraq (3), Jordan (1), Morocco(1), Mozambique (2), Nicaragua (2), Nigeria (3), Pakistan (2), Peru (1), Philippines (1), Rwanda (2), Sierra Leone (1), Somalia (1), Sri Lanka (2), Sudan (1), Turkey (1), Uganda (2), Zaire (1), Zimbabwe (1).

Figures in brackets state the number of civil wars the country experienced.

The hazard of peace ending used data from the following 56 peace episodes

Algeria (2), Argentina (1), Bolivia (1), Burundi (2), Cameroon (1), Chad (1), Columbia (2), Costa Rica (1), Dominican Republic (1), El Salvador (1), Guatemala (3), Greece (1), Kenya (1), India (2), Indonesia (4), Iran (2), Iraq (3), Jordan (1), Malaysia (1), Morocco(2), Mozambique (2), Nicaragua (2), Nigeria (3), Pakistan (2), Paraguay (1), Peru (1), Philippines (2), Rwanda (2), Sri Lanka (1), Sudan (1), Tunisia (1), Uganda (2), Yemen (1), Zaire (1), Zimbabwe (1).

Figures in brackets state the number of peace spells the country experienced.

6.8.2 Variables

The duration of war and peace was obtained from Singer and Small (1994). Only civil wars until 1st January 1993 are listed in this data source. Based on recent Stockholm International Peace Research Institute (SIPRI) yearbooks and newspaper reports we updated the data in the following way to 1st January 1997:

Algeria

After the cancellation of the results of the elections in January 1992 a civil war started in Algeria. It was ongoing by the end of 1996.

Burma (Myanmar)

Singer and Small list Burma as a civil war country at the end of 1992. The level of violence fell below the 1000 battle related deaths in 1994. We see the release of Aung San Suu Kyi on 10 July 1995 as the 'end' of the civil war.

Burundi

Singer and Small list Burundi as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

Columbia

Singer and Small list Columbia as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

India

Singer and Small list India as a civil war country at the end of 1992. The level of violence in Kashmir fell below the 1000 battle related deaths in 1994. The 'end' of the civil war is set at March 1994 when Pakistan made an abortive attempt at censuring India at the UN Human Rights Commission.

Iraq

Singer and Small list Iraq as a civil war country at the end of 1992. The number of battle related deaths fell below 1000 during 94. June 1994 marks the end of the civil war.

Liberia

Singer and Small list Liberia as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

Mozambique

Singer and Small list Mozambique as a civil war country at the end of 1992. However, we treat the Rome Agreement in October 1992 as the end of the war.

Peru

Singer and Small list Sudan as a civil war country at the end of 1992. This war subsided with Guzman's appeal for peace in September 1993.

Philippines

Singer and Small list the Philippines as a civil war country at the end of 1992. The level of violence fell below the 1000 battle related deaths in 1992 and we interpret the election of President Ramos on 11 May 1992 as the end of the civil war.

Rwanda

Singer and Small list Rwanda as a civil war country at the end of 1992. The Rwandan Patriotic Front took control of most of the country by July 1994 which marks the end of the civil war.

Sierra Leone

According to SIPRI the conflict in Sierra Leone began in 1991 and the number of deaths gradually rose. We fix the beginning of the civil war at June 1991. This war was still ongoing by the end of 1996.

Somalia

Singer and Small list Somalia as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

Sri Lanka

Singer and Small list Sri Lanka as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

Sudan

Singer and Small list Sudan as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

Turkey

Singer and Small list Turkey as a civil war country at the end of 1992. This war was still ongoing by the end of 1996.

Yugoslavia

Singer and Small list Yugoslavia as a civil war country at the end of 1992. The cease fire agreed in February 1994 marks the end of the civil war.

Other Variables

Income

Real GDP from Penn World Table Mark 5.6 (RGDPCH).

Ethno-linguistic Fractionalisation Index (ELF)

Index as used by Mauro (1995).

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