Essays on Growth and Environment

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In loving memory of my father,
because the end of life is not the end of love

To Mrs Milena Parea,
the best teacher I could ever have

This work is also dedicated to all the people
who helped me to find my way and added
a bit of "spice" and a "ray of light" to my life

With all my Heart and with all my Love,

Catia

Abstract

This thesis consists of a summary and four self-contained papers.

Paper [I] Following the 1987 report by The World Commission on Environment and Development, the genuine saving has come to play a key role in the context of sustainable development, and the World Bank regularly publishes numbers for genuine saving on a national basis. However, these numbers are typically calculated as if the tax system is non-distortionary. This paper presents an analogue to genuine saving in a second best economy, where the government raises revenue by means of distortionary taxation. We show how the social cost of public debt, which depends on the marginal excess burden, ought to be reflected in the genuine saving. We also illustrate by presenting calculations for Greece, Japan, Portugal, U.K., U.S. and OECD average, showing that the numbers published by the World Bank are likely to be biased and may even give incorrect information as to whether the economy is locally sustainable.

Paper [II] This paper examines the relationships among per capita CO_2 emissions, per capita GDP and international trade based on panel data spanning the period 1960-2008 for 150 countries. A distinction is also made between OECD and Non-OECD countries to capture the differences of this relationship between developed and developing economies. We apply panel unit root and cointegration tests, and estimate a panel error correction model. The results from the error correction model suggest that there are long-term relationships between the variables for the whole sample and for Non-OECD countries. Finally, Granger causality tests show that there is bi-directional short-term causality between per capita GDP and international trade for the whole sample and between per capita GDP and CO_2 emissions for OECD countries.

Paper [III] Fundamental questions in economics are why some regions are richer than others, why their growth rates differ, whether their growth rates tend to converge, and what key factors contribute to explain economic growth. This paper deals with the average income growth, net migration, and changes in unemployment rates at the municipal level in Sweden. The aim is to explore in depth the effects of possible underlying determinants with a particular focus on local policy variables. The analysis is based on a three-equation model. Our results show, among other things, that increases in the local public expenditure and income taxe rate have negative effects on subsequent income income growth. In addition, the results show conditional convergence, i.e. that the average income among the municipal residents tends to grow more rapidly in relatively poor local jurisdictions than in initially "richer" jurisdictions, conditional on the other explanatory variables.

Paper [IV] This paper explores the relationship between income growth and income inequality using data at the municipal level in Sweden for the period 1992-2007. We estimate a fixed effects panel data growth model, where the within-municipality income inequality is one of the explanatory variables. Different inequality measures (Gini coefficient, top income shares, and measures of inequality in the lower and upper part of the income distribution) are examined. We find a positive and significant relationship between income growth and income inequality measured as the Gini coefficient and top income shares, respectively. In addition, while inequality in the upper part of the income distribution is positively associated with the income growth rate, inequality in the lower part of the income distribution seems to be negatively related to the income growth. Our findings also suggest that increased income inequality enhances growth more in municipalities with a high level of average income than in municipalities with a low level of average income.

Keywords: Genuine saving, welfare change, taxation, per capita GDP, per capita CO_2, international trade, net migration, unemployment, growth, inequality.

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Felix, qui potuit rerum cognoscere causas,
atque metus omnis et inexorabile fatum
(Virgilius, "Georgica", II, 490-493)[1]

When I was a child I did not think I would ever go to University, but over the school years, most of my teachers pushed me in that direction. During my education in Italy, I met inspiring teachers, motivating instructors and enthusiastic mentors. Most of them were able to successfully apply to me the Socrates' maieutic art up until my *Laurea* at "La Sapienza" in Rome. My "education journey", after some work experiences, continued abroad; it started some years ago when I moved to Sweden to attend a Master program in Economics. Today, I am even pursuing a Ph.D and I am happy and grateful to many people who helped me to reach this goal.

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[1] Happy the one who has been able to know the causes of things/ and trample on all fears and the inexorable fate./ Virgil, "Georgics", II, 490-493. (Own translation)

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Catia

This thesis consists of a summary and the following four papers:

[I] Aronsson, T., Cialani, C., Löfgren, K-G., (2012). Genuine saving and the social cost of taxation. *Journal of Public Economics*, 96 (1/2), 211–217.

[II] Cialani, C., (2012). CO_2 emissions, GDP and trade: a panel cointegration approach.

[III] Cialani, C. and Lundberg, J., (2013). Growth, migration and unemployment across Swedish municipalities.

[IV] Cialani, C., (2012). Growth and inequality: a study of Swedish municipalities.

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1. Introduction

This thesis consists of four papers. Paper [I] is a theoretical contribution within the area of genuine savings with a focus on a second best economy, where the government raises revenue by means of distortionary taxation. Paper [II] addresses the long-term relationships and short-run causality among per capita GDP, per capita CO_2 and a measure of international trade. Papers [III] and [IV] are empirical contributions to the literature on local economic growth. Paper [III] examines the determinants of average income growth, net migration and unemployment rates in the context of a simultaneous equation system, whereas Paper [IV] analyzes the relationship between income growth and income inequality. Both Paper [III] and Paper [IV] are based on data at the municipal level in Sweden.

This introductory chapter is organized as follows. Section 2 presents and summarizes Paper [I], while Section 3 describes Paper [II]. Section 4 presents the background to the second part of the thesis and summarizes Papers [III] and [IV].

2. Measuring genuine saving

The concept of genuine saving (GS) has, in recent years, become widely accepted as a measure for assessing an economy's local sustainable development[1]. The GS is an indicator of comprehensive net investment i.e., the value of the net investment in all capital stocks of relevance to society. As such, GS does not only reflect the social value of the net investment in physical capital (the measure of the net investment used in conventional national accounting) but also reflects the social value of changes in other capital stocks, such as natural and human capital. An important aspect of GS is that it constitutes an exact measure of welfare change over a short time interval.

A common interpretation is that sustainable development requires welfare to be non-declining over time (Pearce *et al.*, 1989 and Arrow *et al.,* 2003). In economic terms, development is not sustainable if an economy's total stock of capital is not maintained. Genuine saving is a local indicator of sustainable development, where the emphasis on the word "local" is because we are measuring the welfare change over a short time interval; a non-negative number for genuine saving does not imply that the consumption or utility is sustainable over a longer period[2].

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[1] See e.g. Pearce and Atkinson (1993) and Hamilton (1994, 1996). The GS is also implicit (although not explicitly discussed) in Weitzman's (1976) welfare measure. See Aronsson and Löfgren (2010) for an introductory text.
[2] See Asheim (1994) and Pezzey (1994).

As such, the GS indicator has become an important statistic underlying the environmental policy debate with the World Bank regularly publishing estimates for GS[3] on a national basis for over 150 countries since 1970. The conventional approach to measure GS is to add the value of changes in environmental and/or natural capital stocks to the net investment in physical capital, as well as to add the value of net investment in other capital goods such as human capital.

The GS can be derived from a dynamic optimization model where a planner attempts to maximize the present value of social welfare given a set of constraints. Several researchers have pointed out some conceptual issues for the computation of the GS measure (see e.g. Hamilton and Clemens, 1999; World Bank, 1997, 1999) and a few have criticized the method of calculating the resource depletion rates (see e.g. Neumayer, 2000). Dietz and Neumayer (2004) also argue that the GS is sensitive to the method of calculating rents from resource extraction and claim that the World Bank probably overestimates the unsustainability of certain resource-dependent regions. Also, Pillarisetti (2005) argues that the investment in human capital, measured by education expenses, strongly influences the numerical values of the GS and that policy implications based on this measure can be erroneous. Data limitations are also discussed by the World Bank. Thus, the GS can be a misleading indicator of sustainability and we should, therefore, be cautious in interpreting this welfare indicator.

However, the literature on GS has, so far, neglected the fact that tax revenue is typically collected through distortionary taxes. It is, therefore, important to extend the measurement of GS to a second best economy where policy makers face restrictions preventing them from implementing the first best resource allocation. The primary goal of Paper [I] is to present an analogue to GS in a second best economy based on Chamley's (1985) dynamic model, where the government raises revenues through a distortionary tax.

Summary of Paper [I]: Genuine saving and the social cost of taxation

The purpose of this paper is to measure genuine saving when the government raises revenue by means of distortionary taxation. Using a model based on Chamley (1985), we show how the marginal excess burden of taxation contributes to an exact measure of welfare change and, therefore, ought to be reflected in the GS. As such, we modify the conventional measure of GS (calculated as if the tax system is non-distortionary) with an excess of burden that may affect both the sign and magnitude of the welfare change. This means that a second best analogue of the GS may differ in a fundamental way from its first best counterpart.

[3] The World Bank calculations of GS are now called "adjusted net saving".

As a supplement to the theoretical model, we illustrate a application by presenting modified GS numbers for Greece, Japan, Portugal, U.K., U.S. and the OECD average. This example is based on the world-bank numbers supplemented by data on budget deficits and estimates of the marginal excess burden in the literature.

3. Pollution, production and international trade

3.1 Long-term and short-term relationships among GDP, CO_2 and trade

Is free trade good for the environment? The question of whether international trade hurts the environment or not is of great importance, and much attention has been paid to relationships between carbon dioxide (CO_2) emissions, gross domestic product (GDP) and international trade.

Although the long-term relationships and the short-run causality between CO_2 emissions, GDP and international trade have been studied over the past two decades, no consensus regarding the so-called CO_2 emissions-GDP-international trade nexus has yet been reached. More recent studies have addressed the potential simultaneity of increases in pollution, GDP and international trade rather than assuming (possibly erroneously) that trade and GDP are exogenous determinants of pollution (see, e.g., Antweiler, Copeland and Taylor, 2001; Frankel and Rose 2005; Managi, 2006, Managi *et al.*, 2009). Frankel and Rose (2005) use an instrumental variables technique to test for a causal relationship between international trade and environmental pollution (sulfur emissions) by analyzing cross-country data for 1990. More specifically, they address the potential endogeneity of international trade and GDP by applying instrumental variables estimation based on a gravity model of bilateral trade and a growth model derived from neoclassical growth theory. The aim of their work is to study the effect of trade on the environmental pollution for a given level of GDP per capita. They set up a three-equation model: one for the GDP, one for environmental pollution (sulfur emissions) and one for trade openness. Their results show that trade reduces the sulfur emissions. Managi *et al.* (2009) use panel data for sulfur dioxide (SO_2), CO_2 and biochemical oxygen demand (BOD) emissions for 88 countries during the period 1973 – 2000 and BOD data for 83 countries for the period from 1980 to 2000. Managi *et al.* treat trade and income as endogenous and find that the qualitative effect of international trade on pollution varies with the type of pollution as well as between countries. Trade is found to benefit the environment in OECD countries. It has detrimental effects, however, on SO_2 and CO_2 emissions in non-OECD countries, although it contributes to reduce the BOD emissions in these countries.

A recent and emerging line of literature examines the relationship between CO_2 emissions and GDP per capita in a cointegration and short-run causal relationship framework. Dinda and Coondoo

(2006) use a cointegration panel data approach to investigate the relationship between income and per capita CO_2 emissions in 88 countries from 1960 to 1990. They find evidence for cointegration between these variables for the country groups of Africa, Western Europe, Europe and the World as a whole. They also find a bi-directional causality between GDP and emissions for the country group of Central America, Asia and Africa and for the country groups of Europe. This approach also suggests that there is a distinct difference between the short-term and long-term relationships between CO_2 emissions and GDP.

Summary of Paper [II]: CO_2 emissions, GDP and trade: a panel cointegration approach

This paper examines the relationship between the gross domestic product (GDP), carbon dioxide (CO_2) emissions and international trade. We use a dataset for the period 1960 to 2008, based on 150 countries, which are also divided into two subsamples: OECD and Non-OECD countries. This division allows us to capture differences between developed and developing economies. The paper contributes to the literature in several ways; first, by addressing the endogeneity problem arising from the simultaneous determination of CO_2 emissions, GDP and international trade and, second, by using a larger panel dataset in comparison to previous studies (e.g. Dinda and Coondo, 2006; Managi *et al.*, 2009). In addition, the paper addresses panel causality while taking into consideration the heterogeneity in the cross-section units and the non-stationarity of the panel data.

We apply two different classes of cointegration tests: group mean tests and panel tests based on the Error Correction Model (ECM), proposed by Westerlund (2007) and Persyn and Westerlund (2008), to determine the long-run relationship between per capita GDP, per capita CO_2 emissions and a measure of international trade. We test whether the null hypothesis of no error correction can be rejected, either for the whole panel or for at least one cross-unit depending on whether a panel or group mean estimation is carried out. Our findings suggest that when we consider robust p-values, provided through bootstrapping, there is a long-run relationship between per capita CO_2, per capita GDP and international trade for the whole sample and for Non-OECD countries from the panel tests, while the three variables are cointegrated only for Non-OECD countries from the group mean tests. The short-run causality tests show bi-directional short-term causality between per capita GDP and international trade for the whole sample and between per capita GDP and per capita CO_2 emissions for OECD countries. Still, for the OECD sample, our results suggest a causal relationship from per capita GDP and international trade to per capita CO_2 emissions, and from per capita CO_2 emissions and international trade to per capita GDP. For Non-OECD countries, there are two unidirectional relationships, from per capita GDP to international trade and from per capita CO_2 emissions and per capita GDP to international trade.

4. Economic growth at local and regional level

4.1 Determinants of economic growth, migration and unemployment

The question of what factors determine income growth has attracted much attention over the last decades. Many studies on regional growth have taken the hypothesis of unconditional and/or conditional[4] convergence as their point of departure i.e. that poor regions grow faster than rich ones and therefore "catch up" with them. This hypothesis is predicted by neoclassical growth theory as presented by Solow (1956) and Swan (1956). Barro and Sala-i-Martin (1992, 1995) find clear evidence of income convergence between US states, Japanese prefectures and European countries. Using a dataset covering Swedish counties, Persson (1997) finds evidence of unconditional convergence, while Aronsson *et al.* (2001) and Lundberg (2003) find conditional convergence between Swedish counties and municipalities, respectively. Many studies have focused attention on a broader set of possible determinants of regional growth, such as human capital, labor market characteristics and public policy variables (see e.g., Helms, 1985; Barro, 1991; Galeser *et al.*, 1995; Fagerberg *et al.*, 1997). Another example is Aronsson *et al.* (2001), who investigate regional income growth and net migration in Sweden, during the period 1970-1995 and include potential determinants such as local human capital, local labor market characteristics, local public expenditure and investment, intergovernmental grants, demographic characteristics and a measure of political stability and leadership.

Analyses of regional income growth are also closely related to population movements and changes in labor supply. The reason is that income growth may be due to changes in labor supply and/or the composition of the labor force, which makes the parameter estimates of empirical growth models difficult to interpret if the effects of population movements are ignored. Therefore, it is also of importance to include population movements in an analysis of regional income growth. Previous studies of migration have found different economic "opportunity" factors such as the expected wage and the probability of receiving that wage (Treyz *et al.*, 1993; Davies *et al.*, 2001) to be important determinants of migration patterns within the USA.

Regional disparities in average incomes, migration and unemployment rates have been on the Swedish political agenda for decades. One important reason is, of course, that Swedish

[4] Conditional convergence tests focus on the relationship between the income growth rate and the starting position of income, conditional on a set of other explanatory variables.

municipalities are the main providers of welfare services such as child care, primary and secondary education, and care for the elderly, services that are mainly financed by a proportional income tax and through the redistribution system. In particular, the results of Aronsson *et al.* (2001) suggest that future earning possibilities, expressed in terms of initial average income, attract in-migration. Also, the local human capital endowment (measured by the percentage of the population with a university degree) has a positive effect on the subsequent net migration rate. Lundberg (2003) finds local public expenditure and income tax rates to be important determinants of average income growth and net migration at the municipal level. These results are later confirmed by Lundberg (2006).

Consequently, the ability of local governments to provide these services will depend on the growth of the per capita income, and on whether the municipality and the local private sector are successful in attracting labor (net in-migration) and creating jobs (low unemployment).

4.2 Economic growth and inequality

Following Barro (1991), recent literature has addressed the relationship between growth and inequality by incorporating a measure of income inequality as an additional explanatory variable in growth regression models. Based on this approach, and by focusing on cross-country data, an extensive literature has explored how the distribution of income affects the growth rate of an economy's GDP. After Simon Kuznets's (1950)[5] seminal paper in which he argues that there is a trade off-between reducing inequality and increasing economic growth, the mechanisms linking inequality and growth have been extensively addressed in the empirical literature, producing rather conflicting results.

Several studies of the growth-inequality relationship examine a single cross-section of countries and typically find a negative and significant relationship between GDP growth and income inequality (Persson and Tabellini, 1994; Alesina and Perotti, 1996). Other studies, such as Barro (2000) and Forbes (2000), instead report a positive relationship between growth and inequality for relatively rich countries. Similarly, Partridge (1997) finds a positive correlation between GDP growth and the Gini coefficient based on panel data for the US states. Li and Zou (1998) find that the relationship between income inequality and economic growth becomes significantly positive when using panel data (along with appropriate methods for handling such data), while a negative relationship is typically found in single cross-section data. There is also evidence suggesting that the relationship

[5] Kuznets (1955), in his seminal work, suggested the existence of an inverted U-curve between income growth and income inequality known as the Kuznets Curve.

between growth and inequality may vary systematically between countries: Barro (2008) finds that inequality appears to encourage growth in rich countries and slow down growth in poorer countries.

It has been pointed out that, since income statistics may differ significantly between countries due to both the quality and definitions of data, it can be difficult to compare inequality measures between countries (Pardo-Beltrán, 2002). Countries may differ in the "level of democracy, human rights, type of economy, education system etc., which does not make it reasonable to expect that one model holds for all countries" (Nahum, 2005). These issues can be overcome when using data within the same country where many of the institutions are the same.

To our knowledge, there are no previous studies dealing with the relationship between income inequality and growth at the municipal level in Sweden. Previous studies are based on data at the county level (Nahum, 2005) or labor market regions (Rooth and Stenberg, 2012). Nahum (2005) explores the relationship between growth and inequality measured by the Gini coefficient using panel data at the county level from 1960 to 2000. She estimates the effects of inequality measures such as the Gini coefficient, in growth regressions with 1, 3, 5 and 10-year growth periods using fixed effects panel 2SLS estimations and finds a positive effect of inequality on 1 to 5-year economic growth rates (when significant) and no effect on 10-year growth periods. Rooth and Stenberg (2012) analyze the relationship between growth and income inequality based on data for 72 labor market regions in Sweden during the period 1990-2006. The labor market regions were defined using commuting patterns of the labor force provided by Statistics Sweden. Following Voitchovsky (2005), Rooth and Stenberg (2012) investigate whether inequality in different parts of the distribution influences the subsequent economic growth differently. The central hypothesis is that top end inequality encourages growth while bottom end inequality retards growth. Conditional on the Gini coefficient (the measure of overall income inequality), they find a positive relationship between income growth and inequality in the upper part of the income distribution (measured by the ratio between the 90th and 75th income percentiles), while the effect of income inequality in the lower part of the income distribution (measured by the ratio between the 50th and the 10th income percentiles) is insignificant. These results also highlight the potential limitations of studying the impact of inequality on growth based on a single measure of inequality.

Summary of Paper [III]: Growth, migration and unemployment across Swedish municipalities

The purpose of the research described in this paper is to analyze what factors determine the average income growth, migration and unemployment rates at the municipal level in Sweden during the period 1990 – 2007. Although the literature on regional economic growth is quite extensive, this paper contributes to the literature in several ways. In comparison to Fagerberg *et al.* (1996), who examines the effects of the industrial structure on regional economic growth, this paper focuses on the effects of local policy variables, such as taxes and local public expenditures on regional growth as well as by using a broader set of explanatory variables. Moreover, our data cover a longer period (1992-2007) compared to previous studies based on Swedish municipalities (Lundberg, 2003, 2006; Aronsson *et al.*, 2001; Andersson *et al.*, 2007). The paper also extends previous studies based on Swedish data by taking changes in unemployment into consideration. We estimate a three-equation system for the income growth, net migration rate and unemployment rate using three-stage least squares.

The results provide evidence for conditional convergence, both for municipalities in the major city areas and outside the major city areas. These results are in line with the findings in previous studies based on Swedish data (see above). The initial local public expenditures and local taxes are negatively related to the subsequent local income growth. Moreover, the endowment of human capital, measured as the percentage of the population with at least three years of higher education, is positively related to the subsequent net migration and negatively related to the subsequent changes in the unemployment rates.

Summary of Paper [IV]: Growth and inequality: a study of Swedish municipalities

This paper examines whether income inequality among the municipal residents affects income growth, based on data for 283 Swedish municipalities for the period 1992-2007. We consider different measures of income inequality: the Gini coefficient (as a measure of overall inequality) as well as the income shares of the 25%, 15%, 10%, 5% and 1% top income earners. The ratios between the 90th and 75th income percentile (90/75) and 50th and 10th percentile (50/10) are also included to capture inequality at both ends of the income distribution. We also consider interaction effects between the measures of income inequality and the initial income level. The paper contributes to the literature in several ways. First, as far as we know, there are no studies on this topic at municipal level in Sweden. Second, the paper uses a broad set of inequality measures. The baseline model is based on 5-year growth periods, and we test whether the different inequality measures, along with other control variables, affect the average income growth at the municipal level.

The results show that an increase in the Gini coefficient (towards more inequality), increases in the different top income shares, and increased inequality in the upper part of the income distribution all contribute to an increase in the subsequent income growth, while inequality at the bottom of the income distribution is negatively related to income growth. In addition, we find a positive interaction effect between inequality and the initial level of average income, suggesting that increased income inequality stimulates income growth more in rich municipalities than in poor municipalities, *ceteris paribus*.

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Genuine saving and the social cost of taxation☆

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ABSTRACT

Following the 1987 report by The World Commission on Environment and Development, the genuine saving has come to play a key role in the context of sustainable development, and the World Bank regularly publishes numbers for genuine saving on a national basis. However, these numbers are typically calculated as if the tax system is non-distortionary. This paper presents an analogue to genuine saving in a second best economy, where the government raises revenue by means of distortionary taxation. We show how the social cost of public debt, which depends on the marginal excess burden, ought to be reflected in the genuine saving. By presenting calculations for Greece, Japan, Portugal, U.K., U.S. and OECD average, we also show that the numbers published by the World Bank are likely to be biased and may even give incorrect information as to whether the economy is locally sustainable.

1. Introduction

Since the 1970s, a theory of welfare accounting has gradually evolved. One of the basic ideas behind the research in welfare accounting has been to provide a coherent framework for measuring changes in welfare in a dynamic economy, as well as understanding how the current procedures for national accounting ought to be modified with this particular objective in mind.[1] In this paper, we revisit the relationship between capital formation and the subsequent welfare change by presenting a measure of "genuine saving" for a second best economy where the public revenue spent on environmental policy is raised by distortionary taxes. We argue below that such a measure is not only interesting from a theoretical point of view; it has also bearing on statistics of relevance for environmental policy frequently published by the World Bank.

The genuine saving is an indicator of comprehensive net investment, i.e. the value of the net investment in all capital stocks of relevance for society. As such, genuine saving does not only reflect the social value of net investment in physical capital (the measure of net investment used in conventional national accounting); it also reflects the social value of changes in other capital stocks, such as natural and human capital. The remarkable feature with genuine saving is that it constitutes an exact measure of welfare change over a short time interval.[2] Following the 1987 report by The World Commission on Environment and Development, it has also come to play an interesting role as an indicator of sustainable development. The World Commission wrote that development is sustainable if it meets "the needs of the present without compromising the ability of future generations to meet their own needs" (Our Common Future, page 54). One possible interpretation is that sustainable development requires welfare to be non-declining.[3] This suggests, in turn, that the genuine saving is a local indicator of sustainable development, where the emphasis on the word "local" is due to that we are measuring the welfare change over a short time interval.[4] Another interpretation of

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[1] The seminal contribution to the theory of welfare accounting is Weitzman (1976), showing how a welfare-equivalent measure of net national product ought to be defined if the resource allocation is first best. Aronsson (1998, 2008) analyzes the corresponding welfare measurement problem in second best economies, where the public revenue is raised by distortionary taxes. See also the literature reviews by Weitzman (2003) and Aronsson, Löfgren and Backlund (2004).

[2] Although Weitzman (1976) did not attempt to analyze genuine saving, it shows up in the proof of his main result, i.e. we need Weitzman's welfare measure to relate the indicator of welfare change to the genuine saving. Standard references for genuine saving are Pearce and Atkinson (1993) and Hamilton (1994, 1996).

[3] This definition is used in Arrow et al. (2003).

[4] See also Asheim (1994) and Pezzey (1993), who show that a positive value of genuine saving does not give any information as to whether the current level of utility or consumption is sustainable forever.

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sustainable development is that the current (instantaneous) utility level must not fall short of the maximum level that can be sustained forever, in which case non-positive genuine saving indicates that the current utility level faced by the representative consumer exceeds the maximum sustainable level (Pezzey and Toman, 2002; Pezzey, 2004). As such, the genuine saving has become an important statistic underlying the environmental policy debate, and the World Bank regularly publishes numbers for genuine saving on a national basis for a large number of countries.[5]

However, the appropriate procedures for calculating the genuine saving have not received sufficient attention. In fact, the calculations that we have seen either assume that the resource allocation is first best, or that the resource allocation is suboptimal in the sense that society has not reached the best possible outcome given its objective and constraints (due to uninternalized market failures).[6] To our knowledge, there are no studies dealing with the measurement of genuine saving (or an analogue thereof) in economies where the resource allocation is second best optimal; a scenario that will arise if restrictions faced by policy makers prevent them from implementing the first best resource allocation. This gap in the literature is somewhat surprising considering that the revenue raised by the public sector in real world economies typically necessitates distortionary taxes, which are associated with an excess burden that may affect both the sign and magnitude of the welfare change that the economy experiences during a short time interval. Arguably, the principles for measuring genuine saving ought to be modified accordingly; at least if the welfare economic foundation is to be taken seriously. Therefore, the purpose of this paper is to present an analogue to genuine saving in a second best economy, where the government raises revenue through a distortionary tax (instead of a lump-sum tax).

Our study is based on a model developed by Chamley (1985), which is an extension of the Ramsey model in the sense of adding a public sector and assuming that the public revenue is raised by using a linear, yet time-varying, labor income tax. We show that the marginal excess burden of taxation affects the second best analogue to genuine saving via the accumulation of public assets. Finally, we exemplify by adjusting the World Bank numbers for genuine saving and show that neglecting the social costs of taxation (as the World Bank does) may give rise to biased estimates of genuine saving and, in some cases, alter our conclusions as to whether the economy is locally sustainable.

2. The model

The model presented below contains consumers, firms and a government. We start by describing the decision problems faced by agents in the private sector and then continue with the policy problem facing the government. Following much earlier literature, the second best problem will be described as a Stackelberg game, where the government is acting leader and the private agents are followers.

2.1. Consumers and firms

The model developed in this section largely resembles the Ramsey-type models used in earlier literature on welfare accounting with the modification that the public revenue is raised by a labor income tax.[7,8] Following the convention in earlier literature, we assume that the economy is populated by a fixed number of identical consumers normalized to one. The preferences are described by a time-separable utility function. The objective function facing the consumer is represented by the present value of future utility,

$$U(0)=\int_0^\infty u(c(t),z(t),q(t))e^{-\theta t}dt, \tag{1}$$

where c is the consumption of a private good, z leisure and q the quantity of a public good decided upon by the government, while the parameter θ denotes the utility discount rate (i.e. the marginal rate of time preference). The public good is a state variable and may be thought of as public capital that leads to higher environmental quality (e.g., environment-friendly infrastructure, public parks, publicly provided carbon sinks, etc.). This is clearly a somewhat naïve description of environmental quality; by focusing solely on the public sector contribution to such quality, it leaves out a number of vital relationships between production, consumption and damages to the environment. Yet, this simplification is analytically convenient and is of no practical importance for the qualitative relationship between genuine saving and tax distortions, which is the main focus in this paper. As a consequence, we abstract from other aspects of environmental quality. The determination of the public good is discussed below. Leisure is defined as a fixed time endowment, $\bar{l}$, less the hours of work, l. The instantaneous utility function, $u(\cdot)$, is increasing in each argument and strictly concave.

The consumer holds two assets; capital, k, and government bonds, b, which are assumed to be perfect substitutes. If we define $a=k+b$, the asset accumulation equation can be written as

$$\dot{a}(t)=r(t)a(t)+w_n(t)l(t)-c(t) \tag{2}$$

with $a(0)=a_0$, where $w_n(t)=w(t)[1-\tau(t)]$ is the marginal net-of-tax wage rate, in which w is the gross wage rate and τ the tax rate. The variable r is the interest rate. The price of the private consumption good has been normalized to one.

The consumer chooses his/her consumption of the private good, c, and hours of work, l, at each instant to maximize the present value of future utility subject to Eq. (2), the initial condition, and a No Ponzi Game condition (which is a restriction on the present value of the terminal asset). The consumer also treats the factor prices and policy variables at each point in time as exogenous. By using the first order conditions, one can write the demand for the private good and labor supply as functions of the net-of-tax wage rate, the marginal utility of wealth and the public good, respectively,[9]

$$c(t)=c(w_n(t),\phi(t),q(t)) \tag{3}$$

$$l(t)=l(w_n(t),\phi(t),q(t)). \tag{4}$$

The marginal utility of wealth obeys, in turn, the differential equation

$$\dot{\phi}(t)-\theta\phi(t)=-\phi(t)r(t). \tag{5}$$

Finally, by substituting Eqs. (3) and (4) into the instantaneous direct utility function, we obtain the instantaneous indirect utility function

[5] See also Hamilton (2010) for an overview of research on genuine saving.

[6] See, e.g., Aronsson and Löfgren (1998) and Löfgren and Li (2011).

[7] Adding another distortionary tax will not affect the principal findings below. See Chamley (1986) for a dynamic representative agent model with linear taxes on labor income and capital income.

[8] Aronsson (2008) uses a similar model to derive a second best analogue to Weitzman's (1976) welfare measure (i.e. a second best analogue to the comprehensive net national product) when public revenue is collected through distortionary taxes, as well as analyzes the role of public goods in welfare accounting.

[9] Note that the current value Hamiltonian implied by the consumer's decision problem can be written as (if the time-indicator is suppressed)

$$J=u(c,z,q)+\phi\dot{a}$$

where the marginal utility of wealth in current value terms appears as the costate variable attached to wealth. Eqs. (3) and (4) are derived from the first order conditions $u_c(c,z,q)-\phi=0$ and $-u_z(c,z,q)+\phi w_n=0$.

defined conditional on the marginal utility of wealth

$$v(t) = v(w_n(t), \phi(t), q(t)) = u\left(c(w_n(t), \phi(t), q(t)), \bar{l} - l(w_n(t), \phi(t), q(t)), q(t)\right). \quad (6)$$

Turning to the production side, we assume that identical competitive firms use labor and capital to produce a homogenous good under constant returns to scale and normalize the number of firms to one. The production function is given by $f(l(t), k(t))$, and the firm obeys the first order conditions $f_l(l,k) - w = 0$ and $f_k(l,k) - r = 0$.

2.2. The government

The social welfare function coincides with the objective faced by the representative consumer. By using the conditional indirect utility function presented in Eq. (6), the social welfare function at time 0 can be written as

$$V(0) = \int_0^\infty v(w_n(t), \phi(t), q(t)) e^{-\theta t} dt. \quad (7)$$

Turning to the state variables faced by the government, the accumulation equation for the public good is assumed to take the following form:

$$\dot{q}(t) = g(t) - \delta q(t) \quad (8)$$

where $g(t)$ is the contribution to the public good at time t and δ the rate of depreciation. We can think of g as representing public expenditure on investment in environment-friendly infrastructure or abatement of the natural environment. The government uses the income tax to finance the contributions to the public good, although it does not necessarily balance the budget at each instant. If we write the unit tax on labor as $\tau w = w - w_n$, the equation of motion for government bonds is written as

$$\dot{b}(t) = r(t)b(t) + g(t) - [w(t) - w_n(t)]l(t). \quad (9)$$

Finally, by combining Eqs. (2), (9) and the zero profit condition, $f(l,k) - wl - rk = 0$, we can derive the resource constraint

$$\dot{k}(t) = f(l(t), k(t)) - c(t) - g(t). \quad (10)$$

To simplify the notation, we assume that $f(\cdot)$ measures output net of capital depreciation, which means that the left hand side of Eq. (10) represents the net investment in physical capital. Eq. (10) means that output is used for private consumption as well as private and public investment.

The decision problem facing the government will be to choose the tax rate (or net wage rate) and contribution to the public good at each instant to maximize the social welfare function presented in Eq. (7) subject to the state-differential Eqs. (5), (8), (9) and (10), as well as subject to the first order conditions for the private control variables given by Eqs. (3) and (4), and the first order conditions of the firm (which define the gross wage rate and interest rate by the marginal product of labor and capital, respectively). The reason as to why Eq. (5) appears as a state-differential equation in the government's decision problem is that the equation of motion for the private marginal utility of wealth is part of the necessary conditions faced by the consumer and, therefore, a constraint that the optimal tax and expenditure policy must fulfill.[10]

[10] The resource allocation must also obey initial conditions for k and b as well as a No Ponzi Game condition for b. As pointed out by Chamley (1985), the government does not face any explicit constraint on the initial private marginal utility of wealth, $\phi(0)$.

The current value Hamiltonian associated with the public decision problem can be written as (suppressing the time indicator for notational convenience)

$$H = v(w_n, \phi, q) + \lambda \dot{k} + \psi \dot{q} + \mu \dot{b} + \varsigma \dot{\phi} \quad (11)$$

where λ, ψ, μ and ς are the costate variables (measured in current value utility) attached to the state variables in the decision-problem faced by the government, i.e. the stock of physical capital, the environmental public good, the stock of government bonds and the private marginal utility of wealth, respectively. The first order conditions are presented in the Appendix. Here, we use these conditions to derive a measure of welfare change.

3. Measuring genuine saving

The conventional approach to measuring genuine saving is to add the value of changes in environmental and/or natural capital stocks to the net investment in physical capital, as well as adding the value of net investment in other capital goods such as human capital. In our simple model, which abstracts from human capital, this suggests that we should define genuine saving by adding the value of net investment in the environmental public good to the value of net investment in physical capital, i.e. $\lambda \dot{k} + \psi \dot{q}$.[11] We show below that this procedure gives a correct measure of welfare change if the resource allocation is first best, while it does not give a correct measure of welfare change in the second best framework addressed here.

Define the optimal value function at time t as follows:

$$V^0(t) = \int_t^\infty u\left(c^0(s), z^0(s), q^0(s)\right) e^{-\theta(s-t)} ds \quad (12)$$

where $c^0 = c(w_n^0, \phi^0, q^0)$ and $z^0 = \bar{l} - l\left(w_n^0, \phi^0, q^0\right)$ are defined by Eqs. (3) and (4). We use the superindex "0" to denote "second best optimal resource allocation". By totally differentiating the optimal value function represented by Eq. (12) with respect to time, we obtain a measure of welfare change over a short time-interval

$$\frac{dV^0(t)}{dt} = -u\left(c^0(t), z^0(t), q^0(t)\right) + \theta V^0(t). \quad (13)$$

To explore the relationship between the right hand side of Eq. (13) and the measure of genuine saving suggested above, i.e. $\lambda \dot{k} + \psi \dot{q}$, and to be able to relate our study to earlier comparable literature (see the Introduction), we begin by evaluating the welfare change measure in a first best resource allocation. We will then continue with the second best analogue to genuine saving.

3.1. Special case: genuine saving in the first best

In terms of our model, the first best resource allocation constitutes a special case where $\mu(t) = \varsigma(t) = 0$ and $\phi(t) = \lambda(t)$ for all t. Such an allocation would follow if the labor income tax were replaced by a lump-sum tax to finance the contribution to the environmental public good at each instant. By using that $\theta V(t) = H(t)$ at the optimal resource allocation,[12] and if we use the superindex "*" to denote the

[11] Although investment in human capital would affect the exact form of the genuine saving measure, adding human capital to the model would not affect the qualitative results presented below for how the principles of measuring genuine saving ought to be modified in a second best economy by comparison with the corresponding principles in the first best. The welfare measurement problem associated with human capital is addressed by Aronsson and Löfgren (1996).

[12] The optimal control problem is time autonomous, except for the time dependence of the utility discount factor.

first best (to distinguish it from the second best), we obtain the familiar result

$$\frac{dV^*(t)}{dt} = \phi^*(t)\dot{k}^*(t) + \psi^*(t)\dot{q}^*(t). \tag{14}$$

If applied to the model set out above with a two-dimensional capital concept, the right hand side of Eq. (14) is the conventional genuine saving measure. In our model, the genuine saving is given by the sum of the value of net investment in physical and environmental capital. This approach to measure the genuine saving is also consistent with the approach taken by the World Bank; let be that they use a broader capital concept than we do (that also includes human capital).

3.2. Genuine saving in the second best model

Let us now return to the more general second best model set out above. As we show in the Appendix, by applying the same procedure as above, we can derive the following result:

Proposition 1. *The welfare change measure for the second best economy is given by*

$$\frac{dV^0(t)}{dt} = \lambda^0(t)\dot{k}^0(t) + \psi^0(t)\dot{q}^0(t) + \mu^0(t)\dot{b}^0(t) + \varsigma^0(t)\dot{\phi}^0(t). \tag{15}$$

If we follow convention and define genuine saving as $\lambda(t)\dot{k}(t) + \psi(t)\dot{q}(t)$, then the right hand side of Eq. (15) is interpretable as a generalized measure of genuine saving. The generalization follows because the social planner faces two additional state variables here (in addition to k and q); namely, the stock of public debt, b, and the private marginal utility of wealth, ϕ.

The costate variable $\mu(t)$ attached to the government debt at time t is interpretable as the negative of the marginal excess burden at time t; it reflects that increased government debt at present necessitates higher distortionary taxes in the future. If $\mu(t)<0$, as one would normally expect, the intuition is that public debt (asset) accumulation gives rise to a social cost (benefit) due to the distortions generated by the tax system. Therefore, public debt or asset accumulation affects the genuine saving in the second best (which it does not in the first best where $\mu=0$). As pointed out by Chamley (1985), the marginal excess burden measured in real consumption units, $MEB = -\mu(t)/\phi(t) = -\mu(0)/\phi(0)$, is constant over time along the optimal path.[13] Otherwise, it would be possible for the government to reduce the overall welfare cost of taxation by changing its debt policy. We will return to the marginal excess burden below.

The welfare effect of changes in the private marginal utility of wealth, i.e. the fourth term in Eq. (15), is also due to the appearance of distortionary taxation, although for another reason. The tax system distorts the labor supply and private consumption and, therefore, also the path for the private marginal utility of wealth, causing it to differ from the path for the shadow price of physical capital, $\lambda(t)$. The associated welfare cost of this discrepancy is captured by the variable $\varsigma(t)$.[14] To understand why changes in the private marginal utility of wealth affect the welfare change measure, recall that the public decision-problem is formulated in terms of demand functions and an indirect instantaneous utility function, which are defined conditional on the private marginal utility of wealth. If evaluated in the first best, $\partial H/\partial\phi = 0$ (in which case $\varsigma = 0$) because the private cost benefit rule for c and l, respectively, would in that case coincide with the corresponding social cost benefit rule, whereas $\partial H/\partial\phi$ is generally nonzero in the second best due to discrepancies between the private and social cost benefit rules.

If we measure the conventional genuine saving in consumption units, $GS = \left[\lambda\dot{k} + \psi\dot{q}\right]/\phi$, we can rewrite Eq. (15) as follows:

$$\frac{dV^0(t)}{dt} = \phi^0(t)\left[GS^0(t) - MEB^0\dot{b}^0(t) + \frac{\varsigma^0(t)}{\phi^0(t)}\dot{\phi}^0(t)\right]. \tag{16}$$

The expression within square brackets of Eq. (16) is the generalized measure of genuine saving expressed in consumption units. Since $\phi^0(t) > 0$, it follows that welfare increases over the short time-interval $(t, t+dt)$ if, and only if, the expression within square brackets is positive. Eq. (16) constitutes the basis for the application below.

4. Application

In this section, we exemplify by calculating how the marginal value of public debt will modify the numbers for genuine saving published by the World Bank. Our starting point is the World Bank measure of genuine saving, which is defined by subtracting natural resource depletion and damages from emissions (carbon dioxide and particulate matter) from the net investment in physical capital and then adding education expenditures (which serve as a proxy for investment in human capital; let be that the proxy is somewhat misleading, as the connection between such expenditure and the future earnings of the investors is unclear).

To be able to adjust the current numbers for genuine saving, we make two simplifying assumptions: (i) the resource allocation is second best optimal in the sense discussed above, and (ii) the interest rate is constant and equal to the rate of time preference, so $\dot{\phi}(t) = 0$ for all t, in which case the fourth term on the right hand side of Eq. (15) vanishes. The second assumption is needed because there is no way to estimate the value of changes in the private marginal utility of wealth, given the data to which we have access.[15] Therefore, we augment the numbers for genuine saving published by the World Bank by the value of the change in government debt defined as the public deficit times the marginal excess burden measured in consumption units, where the latter is based on estimates in empirical literature.[16]

To be more specific, we will subtract

$$MEB\frac{db(t)}{dt}$$

from the number for genuine saving published by the World Bank. We consider three different numbers for marginal excess burden; 0.1, 0.3 and 0.5, respectively, which are well in line with—although in the lower range of—the estimates summarized by Jacobs (2009).[17] The results are presented in Fig. 1, which contains the numbers published by the World Bank as well as the numbers following the adjustment mentioned above.

Fig. 1 presents the numbers for genuine saving (GS) published by the World Bank, as well as our generalized measure of genuine saving (GGS), for Greece, Japan, Portugal, U.K., U.S. and OECD-average, for the period 1991–2009.[18] All numbers on which the curves are based are

[13] This is seen by solving the equations of motion for $\phi(t)$ and $\mu(t)$ subject to transversality conditions.

[14] In a simplified version of this model without the public good, Chamley (1985) shows that the variable ς is equal to zero at time 0 (due to that there is no initial condition on the private marginal utility of wealth in the second best problem), while it is negative in a steady state.

[15] This is, of course, a restrictive assumption. It may serve as reasonable approximation if the economy is relatively close to a stationary equilibrium.

[16] See Jacobs (2009) for a recent comprehensive literature review.

[17] We realize, of course, that the studies surveyed by Jacobs are typically based on models different from ours. The mean value of marginal excess burden among the studies in his survey is 0.5.

[18] Data for the genuine saving originates from the World Bank and were obtained from the World Development Indicators (collected in the spring of 2011) at http://databank.worldbank.org/ddp/home.do, whereas data for budget surpluses and deficits were collected from the OECD Economic Outlook 87 data base. We used the GNP deflator (UN statistics) to convert the nominal numbers for genuine saving and budget surpluses/deficits into real numbers.

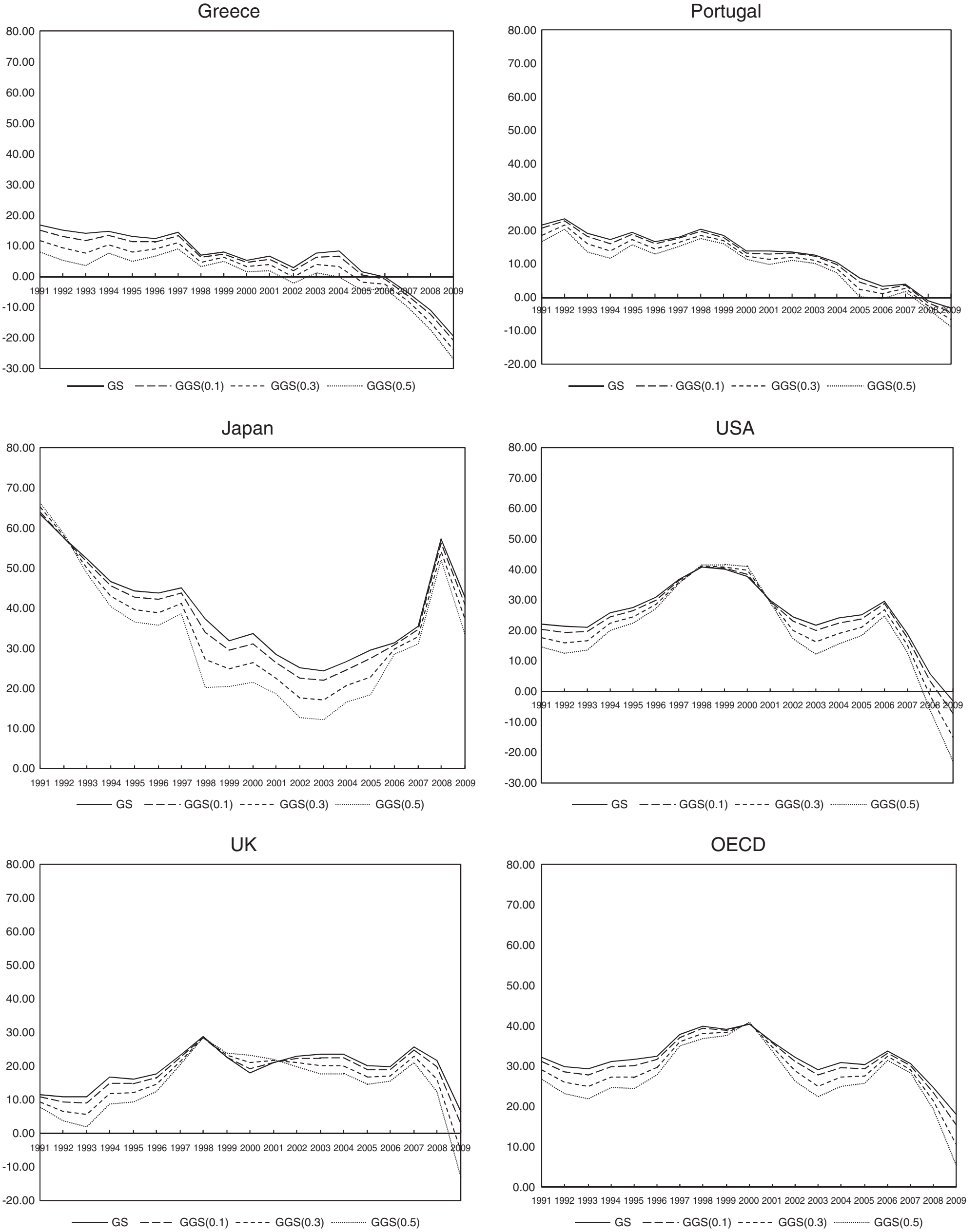

Fig. 1. Genuine saving and generalized genuine saving 1991–2009.

given in hundreds of U.S. dollars in 2009 prices and measured per capita.[19] We make three broad observations from the figure. *First*, neglecting the accumulation of public debt may lead to the wrong conclusion as to whether the economy is locally sustainable. This is the case for the U.S. in 2008, where the conventional genuine saving is positive, whereas the generalized genuine saving is negative if based on the numbers 0.3 and 0.5, respectively, for the marginal excess burden. Similar findings apply for the U.K. in 2009, Greece in 2004–2005 and Portugal in 2006. *Second*, with the highest number for marginal excess burden, i.e. 0.5, which is in line with the empirical evidence referred to above, public debt accumulation may have a considerable effect on the generalized genuine saving. *Third*, and perhaps even more important, since the conventional genuine saving and the budget deficit move together to some extent (e.g., the net investments in physical capital typically fall and budget deficits typically increase during recessions), the conventional genuine saving measure may be a poor indicator as to when the economy is at the risk of becoming locally unsustainable, i.e. the signal that this statistic is designed to give may come several years after which the generalized genuine saving has turned negative.

It is necessary to exercise caution in the interpretation of the results in Fig. 1. One reason is, of course, that the World Bank numbers are uncertain, and it is not always clear that the measure used by the World Bank covers all important aspects of the conventional genuine savings measure—as it ought to be defined in a world without tax distortions—or that all components are measured in the best way possible. The estimates of marginal excess burden are also subject to uncertainty, and the appropriate value to be used may also differ between countries. However, to arrive at an accurate picture of the savings behavior of society, our results suggest, nevertheless, that the savings by the public sector may be of practical relevance when determining whether or not the economy is locally sustainable.

Appendix

For the analysis to be carried out below, it will be more convenient to use the present value Hamiltonian than the current value Hamiltonian. We will, therefore, reformulate Eq. (11) in present value terms, through multiplying by $e^{-\theta t}$, and also write out the constraint functions explicitly. We have

$$\begin{aligned} H_p(t) &= H_p\left(w_n(t), g(t), k(t), q(t), b(t), \phi(t), \lambda_p(t), \psi_p(t), \mu_p(t), \varsigma_p(t), t\right) \\ &= v(w_n(t), \phi(t), q(t))e^{-\theta t} + \lambda_p(t)[f(l(t), k(t)) - c(t) - g(t)] \\ &\quad + \psi_p(t)[g(t) - \delta q(t)] + \mu_p(t)[r(t)b(t) + g(t) - (w(t) - w_n(t))l(t)] \\ &\quad + \varsigma_p(t)\phi(t)[\theta - r(t)] \end{aligned} \tag{A1}$$

where the subindex p attached to the Hamiltonian and costate variables denotes "present value", the factor prices $w(t)$ and $r(t)$ are defined by the marginal products of labor and capital, respectively, while $c(t) = c(w_n(t), \phi(t), q(t))$ and $l(t) = l(w_n(t), \phi(t), q(t))$ according to Eqs. (3) and (4).

The first order conditions for the control variables are (suppressing the time indicator)

$$\frac{\partial H_p}{\partial w_n} = 0 \quad \text{and} \quad \frac{\partial H_p}{\partial g} = 0 \tag{A2}$$

while the equations of motion for the costate variables become

$$\dot{\lambda}_p = -\frac{\partial H_p}{\partial k}, \dot{\psi}_p = -\frac{\partial H_p}{\partial q}, \dot{\mu}_p = -\frac{\partial H_p}{\partial b}, \text{and} \dot{\varsigma}_p = -\frac{\partial H_p}{\partial \phi}. \tag{A3}$$

Derivation of Equation (15)

By totally differentiating Eq. (A1) with respect to time, we obtain

$$\begin{aligned} \frac{dH_p}{dt} = &\frac{\partial H_p}{\partial w_n}\frac{dw_n}{dt} + \frac{\partial H_p}{\partial g}\frac{dg}{dt} + \frac{\partial H_p}{\partial k}\frac{dk}{dt} + \frac{\partial H_p}{\partial q}\frac{dq}{dt} + \frac{\partial H_p}{\partial b}\frac{db}{dt} + \frac{\partial H_p}{\partial \phi}\frac{d\phi}{dt} \\ &+ \frac{\partial H_p}{\partial \lambda_p}\frac{d\lambda_p}{dt} + \frac{\partial H_p}{\partial \psi_p}\frac{d\psi_p}{dt} + \frac{\partial H_p}{\partial \mu_p}\frac{d\mu_p}{dt} + \frac{\partial H_p}{\partial \varsigma_p}\frac{d\varsigma_p}{dt} + \frac{\partial H_p}{\partial t} \end{aligned} \tag{A4}$$

where $\partial H_p/\partial t = -\theta v(w_n, \phi, q)\exp(-\theta t)$, since the direct effect of time is due to the utility discount factor. Now, $\partial H_p/\partial \lambda_p = dk/dt$, $\partial H_p/\partial \psi_p = dq/dt$, $\partial H_p/\partial \mu_p = db/dt$ and $\partial H_p/\partial \varsigma_p = d\phi/dt$ according to the definition of the Hamiltonian. Therefore, by using the first order conditions in Eqs. (A1) and (A2), Eq. (A4) reduces to read

$$\frac{dH_p}{dt} = -\theta v(w_n, \phi, q)e^{-\theta t}. \tag{A5}$$

For notational convenience, we have suppressed the superindex *0* for "second best" in Eq. (A5). Solving Eq. (A5) for $H_p(T)$ gives

$$H_p(T) = H_p(t) - \theta \int_t^T v(w_n(s), \phi(s), q(s))e^{-\theta s} ds. \tag{A6}$$

If T approaches infinity, and by using the transversality condition $\lim_{T\to\infty} H_p(T) = 0$,[20] we obtain

$$H_p(t) = \theta \int_t^\infty v(w_n(s), \phi(s), q(s))e^{-\theta s} ds. \tag{A7}$$

Finally, transforming Eq. (A7) to current value, through multiplying both sides by $e^{\theta t}$, we obtain $\theta V(t) = H(t)$, where $H(t) = H_p(t)e^{\theta t}$ is the current value Hamiltonian and $V(t) = \int_t^\infty v(w_n(s), \phi(s), q(s)) e^{-\theta(s-t)} ds$ the social welfare function.[21] Substitution of $\theta V(t) = H(t)$ into Eq. (13) gives Eq. (15).

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[19] We assumed away population growth in Section 2, because such growth is not important for our qualitative understanding of how the principles for measuring genuine saving in a second best economy differ from the corresponding principles that apply in a first best resource allocation. In practice, population growth adds complications to welfare measurement, since changes in the population affect the welfare change between two subsequent periods (depending on the form of the objective function). We abstract from these complications here.

[20] See Seierstad and Sydsaeter (1987, Chapter 4, page 245).

[21] A result analogous to Eq. (A7) was also derived by Aronsson (1998), although his study focuses on Hamiltonian-based welfare measures and does not address genuine saving.

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II

CO_2 emissions, GDP and trade: a panel cointegration approach

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Abstract

This paper examines the relationships among per capita CO_2 emissions, per capita GDP and international trade based on panel data spanning the period 1960-2008 for 150 countries. A distinction is also made between OECD and Non-OECD countries to capture the differences of this relationship between developed and developing economies. We apply panel unit root and cointegration tests and estimate a panel error correction model. The results from the error correction model suggest that there are long-term relationships between the variables for the whole sample and for Non-OECD countries. Finally, Granger causality tests show that there is bidirectional short-term causality between per capita GDP and international trade for the whole sample and between per capita GDP and CO_2 emissions for OECD countries.

JEL classification: C33, Q28, Q48

Keywords: CO_2 emissions, GDP, international trade, panel data, panel ECM

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1. Introduction

The relationships between economic growth (measured by increases in real GDP per capita) and pollution, as well as between economic growth and international trade, have been analyzed extensively during the last two decades. Also, as countries around the world continue to grow and develop there is increasing interest in elucidating more comprehensively the dynamic relationships among these variables. The purpose of this paper is to estimate the long-term relationships between per capita CO_2 emissions, per capita GDP and international trade; and to examine short-term causal relationships among these variables. To meet these objectives we have analyzed a comprehensive panel data set, and two sub-sets of the data, using the econometric techniques of cointegration and error correction.

There are two well-established lines of research in the literature on this topic. The first originates from studies on environmental economics and is based on joint analysis of GDP and pollution. Much of this work has focused on testing the Environmental Kuznets curve (EKC) hypothesis, according to which there is an inverted U-shaped relationship between pollution and GDP. The EKC hypothesis was proposed and tested in a seminal paper by Grossmann and Krueger (1993). Stern (2004) and Dinda and Coondoo (2006), among others, have reviewed the literature on economic growth and environmental pollution in considerable detail. These reviews demonstrate that no single relationship fits all pollutants for all places and times. However, the existence of an EKC-type relationship has important policy implications. Specifically, policies that stimulate growth (e.g., trade liberalization, economic restructuring, etc.) may reduce environmental pollution in the long run.

The second line of research originates from studies on international economics and is primarily focused on the relationships between international trade on one hand and pollution and GDP growth on the other. Several authors have investigated whether international trade leads to increased pollution as a consequence of increased production or income (e.g., Copeland and Taylor, 1994; Rodríguez and Rodrik, 1999; Frankel and Romer, 1999; Frankel and Rose, 2002). These studies indicate that international trade can affect the environment, even if the empirical relationships between trade, GDP and different types of pollution are not clear-cut. For instance, openness to trade can have positive or negative effects on the environment (Grossman and Krueger, 1993), because the overall effect is due to the combined impact of changes in industrial composition, increasing GDP, and increasing demand for environmental quality. Furthermore, there is an extensive body of literature on the relationship between economic growth and international trade (see, for example, the surveys by Edwards, 1998; Giles and Williams, 2000a, 2000b; and Lewer and

Van den Berg, 2003). Much of this research deals with the link between exports and GDP by testing the export-led growth and growth-led export hypotheses. Different studies have yielded substantially divergent results, making it difficult to draw unambiguous conclusions. More recent studies have addressed the potential simultaneity of increases in pollution, GDP (or national income) and international trade rather than assuming (possibly erroneously) that trade and GDP are exogenous determinants of pollution (see Antweiler *et al.*, 2001; Frankel and Rose, 2005; Managi, 2006; Managi *et al.*, 2009). Frankel and Rose (2005) used an instrumental variables technique to test for a causal relationship between international trade and environmental pollution by analyzing cross-country data for 1990. The central focus of their work was the effect of trade on the environment for a given level of GDP per capita. They derived three equations: one for GDP, one for environmental pollution (specifically sulfur emissions) and one for trade. They also examined the endogeneity of trade openness, which was included as an explanatory variable in both the GDP and environmental quality equations, by introducing a gravity model of bilateral trade as a research instrument. The three derived equations were then used to test the validity of a proposed causal relationship between international trade and environmental pollution. Their results show that trade reduces sulfur dioxide emissions.

Some of the studies from both lines of research have focused on the relationship between GDP and the environment (e.g., the EKC) or between GDP and trade, while other authors such as Frankel and Rose (2005) and Managi (2006) and Managi *et al.* (2009) have studied the nexus among CO_2 emissions, GDP growth and international trade using a single unified model to explicitly describe the endogeneity of GDP and trade.

Why is it interesting to study the nexus among GDP, international trade and CO_2 pollution? Much attention has been paid to global environmental problems: in particular the relationship between CO_2 emissions and trade liberalization policies. The debate focuses on two different but related issues (Huang and Labys, 2001). The first, following the agenda of the Kyoto Protocol, is the rising trend in carbon emissions. One of the most important challenges for environmental policy in the near future will be to reduce these emissions. Hence, an understanding of the relationships between CO_2 emissions, GDP and international trade is essential for formulating effective public policy. The second major issue is trade openness, which probably promotes GDP growth but may also increase pollution. The ongoing globalization of the world's economy is increasing the volume of international trade, and this has further contributed to the growing interest in the relationships between international trade, economic growth and environmental pollution.

We have examined the relationships between CO_2 emissions, GDP and international trade by using three time-series econometric techniques—unit root testing, cointegration and the related Error Correction (EC) model—to analyze a panel data set. One of the key objectives of the present paper is to determine whether the time series for CO_2 emissions, GDP and international trade follow similar temporal trends. In addition, the directions of short-run causality among these three variables are examined. The analyzed data set consists of a data panel covering 150 countries for the 1960–2008 period. Separate estimates are presented for all countries, OECD countries and Non-OECD countries. The sample was split into these two groups of countries because most developing countries, which are heavily represented among the Non-OECD economies, are not signatories of the Kyoto Protocol. Consequently, the relationships that this paper attempts to capture are likely to differ substantially between developed and developing economies. Moreover, in recent decades many poor countries have experienced rapid economic development after adopting liberal economic policies (Akyüz and Gore, 2001).

Our analysis is based on the panel EC and cointegration approach recently proposed by Westerlund (2007) to test whether CO_2, GDP and a common measure of international trade are cointegrated, i.e., whether there is a stationary linear combination of the random variables CO_2, GDP and international trade. The heterogeneous panel unit root test developed by Im, Pesaran and Shin (2003) is used to check for stationarity. This paper thus fills a gap in the literature by using a dynamic panel error correction model to study the causal linkages among all three variables. In our framework, the per capita GDP, the measure of international trade, and per capita CO_2 emissions are treated as three potentially simultaneous variables, and the issue of short-run causality is addressed through a series of regressions where each variable is regressed against the other two.

The paper contributes to the literature in several ways. First, our work uses a larger dataset than previous studies on similar topics that have used a panel approach of any kind. Dinda and Coondoo (2006) used a panel data–based cointegration approach to study incomes and emissions in 83 countries over 30 years, while Managi *et al.* (2009) used panel data for SO_2 and CO_2 emissions of 88 countries over 27 years, and Biological Oxygen Demand (organic pollutant) emissions of 83 countries over 20 years. Our data set includes 150 countries as a full sample, 30 OECD countries and 120 Non-OECD countries, over a period of 48 years. Second, the cointegration approach allows us to address the endogeneity problem that arises from the simultaneous determination of CO_2 emissions, GDP and international trade. This has been one of the most extensively discussed issues in previous publications on trade and the environment (e.g. Frankel and Rose, 2005; Managi, 2006; Managi *et al.*, 2009). Third, most empirical studies focus on either the relationship between pollution and GDP or that between GDP and international trade. Very few (notable exceptions are the works of Managi, 2006, and Managi *et al.*, 2009) are based on panel data, primarily because of the lack of data on pollutant levels over longer periods of time. In contrast, our approach enables us to model the determination of CO_2 emissions, GDP and international trade simultaneously, and examine how these variables change over time in both the short and long run. Fourth, our panel causality tests take into consideration the heterogeneity in the cross-section units and the non-stationary aspects of the panel structure of our data, both of which are neglected in most panel causality studies in this field.

The rest of the paper is organized as follows: Section 2 presents the data and the empirical methodology. Section 3 illustrates and discusses the empirical results. Finally, Section 4 presents and discusses our conclusions.

2. Empirical Framework

2.1 Data sources and variables

As mentioned in the introduction, the full sample consists of data for 150 countries covering the 1960–2008 period[1]. Separate estimates were prepared for two groups of countries: the OECD nations (30 countries) and the Non-OECD nations (120 countries). The basic country-level data—i.e. per capita real GDP (PPP) and information about international trade (exports and imports)—were obtained from the Penn World Table (Mark 7.0). In the analysis below, the per capita GDP is expressed in US$ measured in real 2005 PPP-adjusted dollars and converted to log form, while the indicator of international trade is defined as exports plus imports divided by GDP, i.e. the total volume of trade as a proportion of GDP. The corresponding country-level annual data on per capita CO_2 emissions, expressed in metric tons, were obtained from the Tables of National CO_2 Emissions prepared by the Carbon Dioxide Information Analysis Center, Environmental Science Division, Oak Ridge National Laboratory, United States.

The standard summary statistics of our data are available in Appendix A, while the list of countries included in the analysis can be found in Appendix C. As can be seen from Table A1, the mean per capita CO_2 emissions are higher for OECD than for Non-OECD countries. In addition, Non-OECD countries exhibit the greatest range (distance between Max and Min) in metric tons of CO_2 released per capita. A similar trend is observed for per capita GDP, with the mean being greater for the OECD countries than for the Non-OECD countries. In addition, per capita CO_2 emissions, per

[1] We omitted 16 countries for which we had insufficient historical data on international trade and CO_2 emissions.

capita GDP and volume of international trade are all more variable (as judged by the corresponding standard deviations) between countries than within countries.

With respect to international trade, Table A1 shows that Non-OECD countries are more open to trade than OECD countries (as indicated by the ratio of total trade to GDP). It should be noted that the Non-OECD sample includes some high- and medium-income countries according to World Bank classifications.

Hereafter, log values of real GDP per capita[2] are denoted Y, per capita CO_2 emissions E and the measure of international trade T.

2.2 Econometric Technique

As indicated in the introduction, this paper examines the relationships among E, Y and T. To address the stationarity properties of the time series, a panel data unit root test is performed to determine whether or not the observed country-specific time series for Y, E and T exhibit stochastic trends. Next, cointegration analysis is performed to examine whether the variables are cointegrated (i.e. whether there are stable long-term equilibrium relationships among them). Finally, an Error Correction Model (ECM) is estimated, to test the short-term causality relationships among E, Y and T.

Panel unit root test

As a first step, we must determine the order of integration of the three series in our data. Testing for unit root is performed using the panel unit root test of Im, Pesaran and Shin (2003; hereafter the IPS test), which is appropriate for balanced panels:

$$\Delta x_{it} = \alpha_i + \tau_i t + \rho_i x_{i,t-1} + \sum_{j=1}^{h_i} \beta_{ij} \Delta x_{i,t-j} + \varepsilon_{it} \tag{1}$$

for $i = 1,2,...,N$ and $t = 1,2,...,T$, where $x = E,Y,T$, i and t denote cross-sectional unit and time, respectively, α_i, τ_i and β_i are parameters, ρ_i, is the autoregressive root and h_i is the number of lags and $\varepsilon_{i,t}$ is the error term. The null hypothesis of this test means that all series in the panel are non-stationary processes, so $H_0 : \rho_i = 0, \forall i,$ and the corresponding alternative hypothesis is that some (but not necessarily all) of the individual series in the panel are stationary, i.e. $H_1 : \rho_i < 0$ for at

[2] We take the log of GDP for scale reasons and to facilitate interpretation of the coefficients.

least one i. This test is based on the Augmented Dickey-Fuller (ADF) testing approach and defines the *t*-bar statistic, $\bar{t}$, as a simple average of the individual ADF statistics for all i (denoted as $t_{\rho i}$):

$$\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t_{\rho i}$$

where $t_{\rho i}$ is the individual *t*-statistic for testing $H_0 : \rho_i = 0, \forall i.$ It is shown that, given *N* as $T \to \infty$, $t_{\rho i}$ weakly converges to t_{iT} [3].

In order to have a standardization of the $\bar{t}$ statistics, the IPS test assumes that the individual t_{iT} are *iid* and have finite mean and variance. Im, Pesaran and Shin (2003) have proposed the following panel unit root test statistic, $W_{[t-bar]}$, using the means and the variances of t_{iT}, evaluated under the null $\rho_i = 0$, which is applicable to heterogeneous cross-sectional panels:

$$W_{[t-bar]} = \frac{\sqrt{N}\left(\bar{t} - 1/N\sum_{i=1}^{N} E\left[t_{iT} \middle| \rho_i = 0\right]\right)}{\sqrt{Var\left[t_{iT} \middle| \rho_i = 0\right]}}$$

where $E\left[t_{iT} \middle| \rho_i = 0\right]$ and $Var\left[t_{iT} \middle| \rho_i = 0\right]$ denote, respectively, the moments of mean and variance computed by using Monte Carlo simulations for different values of T and ρ_i's by Im, Pesaran and Shin (2003). The statistic $W_{[t-bar]}$ approaches a standard normal distribution as N and $T \to \infty$.

Error Correction based Panel Cointegration tests

As a second step, we apply the panel cointegration tests developed by Westerlund (2007) and Persyn and Westerlund (2008). The rationale here is to test for the absence of cointegration by determining whether Error Correction exists for individual panel members or for the panel as a whole.

Consider the Error Correction Models described by equations (2), (3) and (4), in which all variables in levels are assumed to be integrated of order one, *I*(1):

[3] For more details how to compute t_{iT}, see Hamilton (1994), chap. 17, p. 478, and Im, Pesaran and Shin (2003).

$$\Delta E_{i,t} = \alpha_i^E + \lambda_i^E(E_{i,t-1} - \beta_i^E Y_{i,t-1} - \gamma_i^E T_{i,t-1}) + \sum_{j=1}^{n} \theta_{i,j}^E \Delta E_{i,t-j} + \sum_{j=1}^{p} \phi_{i,j}^E \Delta T_{+i,t-j} + \sum_{j=1}^{m} \delta_{i,j}^E \Delta Y_{i,t-j} + u_{i,t} \quad (2)$$

$$\Delta Y_{i,t} = \alpha_i^Y + \lambda_i^Y(Y_{i,t-1} - \beta_i^Y E_{i,t-1} - \gamma_i^Y T_{i,t-1}) + \sum_{j=1}^{n} \delta_{i,j}^Y \Delta Y_{i,t-j} + \sum_{j=1}^{m} \theta_{i,j}^Y \Delta E_{i,t-j} + \sum_{j=1}^{p} \phi_{i,j}^Y \Delta T_{+i,t-j} + \varepsilon_{i,t} \quad (3)$$

$$\Delta T_{i,t} = \alpha_i^T + \lambda_i^T(T_{i,t-1} - \beta_i^T Y_{i,t-1} - \gamma_i^T E_{i,t-1}) + \sum_{j=1}^{p} \phi_{i,j}^T \Delta T_{i,t-j} + \sum_{j=1}^{m} \delta_{i,j}^T \Delta Y_{i,t-j} + \sum_{j=1}^{n} \theta_{i,j}^T \Delta E_{+i,t-j} + e_{i,t} \quad (4)$$

Here, the parameters $\lambda_i^k,\ k \in \{E,Y,T\}$ are the parameters of the *Error Correction* (EC) term and provide estimates of the speed of error correction towards the long-run equilibrium for country *i*, while $\varepsilon_{i,t}$, $u_{i,t}$ and $e_{i,t}$ are white noise random disturbances.

We focus on E and its relation to Y and T; therefore, equation (2) is our primary equation of interest. Equations (3) and (4) can potentially be ignored if Y and T can be treated as weakly exogenous, and the validity of this assumption can be tested by performing regressions with $\Delta Y_{i,t}$ and $\Delta T_{i,t}$ as dependent variables.

Two different classes of tests can be used to evaluate the null hypothesis of no cointegration and the alternative hypothesis: group-mean tests and panel tests. Westerlund (2007) developed four panel cointegration test statistics (G_a, G_t, P_a and P_t)[4] based on the Error Correction Model (ECM). The group-mean tests are based on weighted sums of λ_i^k estimated for individual countries, whereas the panel tests are based on an estimate of λ^k for the panel as a whole. These four test statistics are normally distributed. The two tests (G_t, P_t) are computed with the standard errors of λ_i^k estimated in a standard way, while the other statistics (G_a, P_a) are based on Newey and West (1994) standard errors, adjusted for heteroscedasticity and autocorrelations. By applying an Error Correction Model in which all variables are assumed to be *I*(1), the tests proposed by Westerlund (2007) examine

[4] $G_t = \frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\lambda}_i^k}{s.e.(\hat{\lambda}_i^k)}$, $G_a = \frac{1}{N}\sum_{i=1}^{N}\frac{T\hat{\lambda}_i^k}{\hat{\lambda}_i^k(1)}$; $\lambda_i^k(1) = \hat{\omega}_{ui}/\hat{\omega}_{xi}$ where $\hat{\omega}_{ui}$ and $\hat{\omega}_{xi}$ are the usual Newey and West (1994) standard error corresponding to the long-run variance estimators $\hat{\omega}_{xi}^2 = \frac{1}{T-1}\sum_{j=-M_i}^{M_i}\left(1 - \frac{j}{M+1_i}\right)\sum_{t=j+1}^{T}\Delta x_{it}\Delta x_{it-j}$ where M_i is a bandwidth parameter that determines how many covariances to estimate in the kernel. $\hat{\omega}_{ui}$ may be obtained as above using kernel estimation with Δx_{it} replaced by $\hat{v}_{it} = \sum_{j=1}^{m}\hat{\delta}_{i,j}^E + \hat{u}_{i,t}$ for (2) and so on for (3) and (4). $P_t = \frac{\hat{\lambda}_k}{s.e.(\hat{\lambda}_k)}$, $P_a = T\hat{\lambda}_k$.

whether cointegration is present or not by determining whether error correction is present for individual panel members and for the panel as a whole.

If $\lambda_i^k < 0,$ then there is an error correction, which implies that $Y_{i,t}$ and $E_{i,t}$ and $T_{i,t}$ are cointegrated, whereas if $\lambda_i^k = 0$ there is no error correction and thus no cointegration. Thus, the null hypothesis of no cointegration for the group-mean tests (G_a and G_t test statistics) is as follows: $H_0^G : \lambda_i^k = 0$ for all i, which is tested against $H_1^G : \lambda_i^k < 0$ for at least one i. In other words, in the two group-mean-based tests, the alternative hypothesis is that there is cointegration in at least one cross-sectional unit. Therefore, the adjustment coefficient λ_i^k may be heterogeneous across the cross-section units. Rejection of H_0^G should therefore be taken as evidence of cointegration in at least one of the cross-sectional units. The panel tests (P_a and P_t test statistics) instead assume that $\lambda_i^k = \lambda^k$ for all i, so the alternative hypothesis is that adjustment to equilibrium is homogenous across cross-section units. Then, we test $H_0^P : \lambda^k = 0$ against $H_1^P : \lambda^k < 0.$ Rejection of H_0^P should therefore be taken as evidence of cointegration for the panel as a whole.

The tests are very flexible and allow for an almost completely heterogeneous specification of both the long-run and short-run parts of the error correction model. The series are allowed to be of unequal length. If cross-sectional units are suspected to be correlated, robust critical values can be obtained through bootstrapping of the test statistics.

We are mainly interested in the long-run behavior of our model, so the next step is to determine the coefficients of the conditional long-run relationships between *E*, *Y* and *T* when the short-run terms are set to zero. The long-run coefficients can be easily derived from the following long-run equation, obtained from the reduced form of (2) when the terms representing short-run changes are $\Delta E = \Delta T = \Delta Y = 0$, as follows:

$$E_{i,t} = -\frac{\alpha_i^E}{\lambda_i^E} + \beta_i^E Y_{i,t} + \gamma_i^E T_{i,t}$$

Finally, we also test for short-run causality. This implies testing the significance of the coefficients of the lagged difference of the variables (using the Wald restriction test) for equations (2), (3) and (4). The putative causality of individual relationships is tested by checking the significance of the *t*-statistic for the coefficient of the lagged variable, while the joint causality is tested as follows.

We can test the null hypotheses that the other two variables are not sources of short-run causation of *E*, Y and T by testing whether $H_0 : \varphi_{i,j}^E = \delta_{i,j}^E = 0 \forall i,$ $H_0 : \theta_{i,j}^Y = \phi_{i,j}^Y = 0 \forall i$ and

$H_0 : \theta^T_{i,j} = \delta^T_{i,j} = 0 \forall i$ (Eqs. 2, 3 and 4), respectively and if these null hypotheses are rejected, we will have bi-directional causality.

3. Results and discussion

The panel unit root test results for E, Y and T over the full sample are summarized in Table 1. The decision of whether or not to reject the null hypothesis of unit root for the panel as whole is based on the $W_{[t-bar]}$ statistic.

Table 1: Im-Pesaran-Shin Test for Unit Root in Panels for the full sample

	LEVELS		FIRST DIFFERENCES	
Variable	**Constant**	**Constant and Trend**	**Constant**	**Constant and Trend**
	$W_{[t-bar]}$ **Statistic**			
E	-3.237***	-1.033	-24.923***	-19.998***
Y	21.922	10.907	-22.094***	-18.305***
T	6.054	4.588	-33.093***	-30.063***

Note: *** indicates significance at the *P*<0.01 level.

We were not able to reject the null unit root hypothesis for the Y and T series when expressed in level form. However, E is stationary without a trend term. When using the first differences, the null of unit roots is strongly rejected at the *P*<0.01 significance level for all three series, implying that the series are *I*(1). This finding is confirmed by all tests employed for all three country samples examined: i.e. the full sample and both the OECD and Non-OECD sub-samples, although the corresponding values are not presented herein.

We proceed by testing whether Y, E and T are cointegrated (see Appendix B for the specifications used in the four cointegration tests). We adopt the Westerlund-based panel cointegration tests using a single lag and lead, $h_i = q_i = 1$. The lead and lag orders were selected based on the minimum AIC (Akaike's Information Criterion). We perform cointegration tests with both a constant and a trend, no constant or trend, and with a constant but no trend. We also consider the robust *P*-values obtained after bootstrapping using 800 replicates after testing for cross-sectional dependence among residuals.

Results obtained from the model with a constant but no trend suggest that there is no cointegration for Y and T (Table 2; see Table B1 in Appendix B for results from the other cointegration tests). However, as can be seen in Table 2, our results for the whole sample—i.e. from the panel cointegration tests—indicate that there is a long-run cointegrating relationship for E among the series under consideration, based on equation (2). The P_t and P_a statistics indicate that the null hypothesis of no cointegration for E should be rejected at the *P*<0.01 level. The other models (neither constant nor trend, and both a constant and trend) also indicate that the null hypothesis of no cointegration for E should be rejected at the *P*<0.01 level. The robust *P*-values indicate that the null hypothesis of no cointegration should be rejected at the *P*<0.05 level for the full sample and the *P*<0.01 level for Non-OECD countries.

As can be seen from the *P*-values, for the income equation (Y) the null hypothesis of no cointegration cannot be rejected for either the full sample or the OECD sample. However, the P_t and P_a values indicate that the null hypothesis of no cointegration (and hence no stationary equilibrium relationship among the variables) should be rejected at *P*<0.01 for the Non-OECD sample. At the same time, the robust *P*-values indicate that we cannot reject the null hypothesis of no cointegration for either the full sample or the OECD and Non-OECD countries.

Table 2: Results of the Westerlund-based Panel Cointegration tests

		with Constant but No Trend											
Model	**Test**	**Full sample**				**OECD**				**Non-OECD**			
		value of Test	**z-value**	**P-value**	**Robust P-value**	**value of Test**	**z-value**	**P-value**	**Robust P-value**	**value of Test**	**z-value**	**P-value**	**Robust P-value**
Y	G_t	-1.706	4.379	1.000	0.998	-1.585	2.676	0.996	0.954	-1.720	3.746	0.996	0.908
	G_a	-5.885	6.326	1.000	0.995	-5.508	3.158	0.999	0.934	-5.966	5.516	1.000	0.946
	P_t	-18.152	2.740	0.997	0.789	-7.420	1.901	0.971	0.745	-64.092	-43.894	0.000	0.858
	P_a	-4.919	2.073	0.981	0.523	-4.263	1.569	0.942	0.614	-18.629	-25.007	0.000	0.759
E	G_t	-2.01	0.329	0.629	0.741	-1.916	0.706	0.760	0.741	-2.032	0.040	0.516	0.002
	G_a	-7.701	2.78	0.997	0.171	-5.147	3.473	1.000	0.984	-8.336	1.377	0.916	0.024
	P_t	-27.016	-5.845	0.000	0.030	-8.020	1.320	0.907	0.735	-24.819	-5.862	0.000	0.000
	P_a	-8.023	-4.728	0.000	0.000	-4.098	1.731	0.958	0.794	-8.401	-4.968	0.000	0.000
T	G_t	-1.693	4.550	1.000	0.999	-0.791	7.400	1.000	1.000	-1.545	5.827	1.000	1.000
	G_a	-5.686	6.715	1.000	0.958	-1.311	6.824	1.000	1.000	-4.695	7.737	1.000	1.000
	P_t	-20.841	0.135	0.554	0.121	-2.225	6.932	1.000	0.980	-17.871	0.867	0.807	0.890
	P_a	-6.051	-0.407	0.342	0.203	-1.262	4.509	1.000	0.966	-5.495	0.726	0.766	0.818

Note: We then used xtwest to test for cointegration, using the AIC to choose the optimal lag and lead lengths for each series and with the Bartlett kernel window width set to $4(T/100)^{2/9} \approx 3$.

For the trade equation (T), there is cointegration across the panel as a whole when the model is estimated without constant and trend terms. However, the addition of either a constant alone or a constant and a trend term makes all of the test statistics non-significant for all of the samples. Thus, the null hypothesis of no cointegration in the trade equation cannot be rejected for the model with either a constant or both constant and trend terms.

Because of differences in their construction, "group-mean" and "panel" tests can give different results, and the G_a and G_t test statistics do not indicate that the null hypothesis of no cointegration can be rejected, even at $P<0.10$ (except for E in the Non-OECD countries, for which the robust *P*-values of the G_a and G_t test statistics indicate that the null hypothesis can be rejected at the $P<0.05$ level).

Caution is required when interpreting the results of our tests for the emission equation. Given the definitions used, one would expect the group-mean tests to reject the null hypothesis more often than the panel tests (because at least one series is cointegrating in the former case, which might not necessarily show up in the latter test), and not the opposite. When analyzing a small data set, such as that used here (T=48), the results of the two tests should be interpreted carefully[5]. As a consequence, for our data it seems that panel tests are probably more appropriate than group-mean tests[6].

The economic implication of the existence of cointegration is that there is a stable equilibrium long-run relationship among the variables E, Y and T. Table 2 provides evidence of cointegration in the emissions equation for both the full sample and Non-OECD countries. However, the other models suggest that there is no cointegration of Y, except for some evidence of cointegration for Y based on the *P*-value obtained from the panel tests for Non-OECD countries. Thus, results based on the income equation should be interpreted with caution.

A further consideration is that our results are somewhat mixed, especially the robust *P*-values. For both the full sample and Non-OECD countries, only the panel tests suggest there are long-run relationships among E, Y and T. When we account for cross-sectional dependence using the bootstrap approach, we get somewhat different results. For both the full sample and Non-OECD countries, cointegration is still confirmed by the panel tests; however, for Non-OECD countries the group-mean tests also indicate that the no cointegration hypothesis should be rejected.

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[5] The group-mean and panel tests are constructed in different ways and can therefore give different results. They require large N and large T datasets. These tests are also very sensitive to the specific choice of parameters such as lag and lead lengths, and the kernel width.

[6] We also estimated the group-mean error-correction model, averaging coefficients of the error-correction equation over all cross-sectional units, together with the implied long-run relationship. However, these results are not reported here because the long-run coefficients were not significant.

Overall, the primary model used in this study suggests that there are long-run relationships among E, Y and T for the whole sample as a panel and for Non-OECD countries, both as a panel and as individual panel members.

Error Correction Model estimates

Given the evidence of panel cointegration, the long-run relationships among E, Y and T can be further estimated by applying the estimator of Westerlund (2007). Therefore, we estimate equations (2), (3) and (4) of the ECM, reparameterized based on panel data. Table 3 reports the findings for the three specifications for comprehensiveness, although our focus is on E (Eq. 2).

We approach the interpretation of the regression results presented in Table 3 from the point of view of short-run fluctuations around a long-run equilibrium relationship. In Table 4 we report the results for the long-run relationships of E, Y and T, while Table 5 presents results of the test of the short-run causality relationships. In Table 3, all of the estimated adjustment parameters (i.e., the coefficients of the EC term) are statistically significant and have the expected negative sign, except those for the OECD countries when T is taken as the dependent variable. This result is consistent with the findings reported by Dinda and Coondoo (2006), of negative coefficients for Africa, Central America, America as whole, Eastern Europe, Europe as a whole and the world.

In the equation for E, we find that λ^E is negative for each of the three country groups. This implies that if $E_{t-1} > \beta^E Y_{t-1} + \gamma^E T_{t-1}$, the EC term induces a negative change in E back toward the long-run equilibrium. We obtain larger absolute values for the Non-OECD countries (0.142) and the full sample (0.130) than for the OECD countries (0.055). This implies that a much longer time will be required for equilibrium to be restored following any deviation from the long-run equilibrium in the OECD countries than in the Non-OECD countries. Therefore, this empirical evidence suggests structural divergences between the OECD and Non-OECD countries in the speed of adjustment towards long-run equilibrium.

Table 3: Results of the ECM Estimates

Regressors	FULL SAMPLE			OECD			Non -OECD		
	ΔY	ΔE	ΔT	ΔY	ΔE	ΔT	ΔY	ΔE	ΔT
Constant	0.249*** (12.76)	-0.309*** (-3.51)	-16.44*** (-5.68)	0.310*** (9.44)	-0.548** (-2.69)	-8.921** (-2.92)	0.465*** (17.86)	-0.410*** (-4.17)	-8.113* (-2.44)
$Y_{(t-1)}$	-0.030*** (-12.06)	0.053*** (4.71)	2.922*** (7.92)	-0.032*** (-8.46)	0.075** (3.24)	0.975** (2.80)	-0.060*** (-17.34)	0.062*** (4.77)	2.126*** (4.86)
$E_{(t-1)}$	0.003 (1.88)	-0.130*** (-20.89)	-0.138 (-0.63)	0.000240 (0.15)	-0.055*** (-6.24)	-0.320* (-2.36)	0.00800*** (3.73)	-0.142*** (-19.49)	0.153 (0.58)
$T_{(t-1)}$	0.021*** (5.57)	0.002 (0.15)	-0.093*** (-17.68)	0.039*** (5.64)	-0.173*** (-4.23)	0.022*** (3.57)	0.028*** (5.38)	0.019 (0.99)	-0.104*** (-17.16)
$\Delta Y_{(t-1)}$	0.123*** (10.45)	0.087 (1.64)	7.135*** (4.11)	0.230*** (9.06)	0.191 (1.25)	-7.008** (-3.10)	-0.141*** (-12.64)	0.060 (1.45)	-0.056 (-0.04)
$\Delta T_{(t-1)}$	0.020* (2.40)	0.017 (0.47)	-0.088*** (-7.31)	0.077* (2.32)	-0.323 (-1.66)	0.104*** (3.58)	-0.011 (-1.00)	0.016 (0.29)	-0.098*** (-7.31)
$\Delta E_{(t-1)}$	0.002 (0.89)	-0.057*** (-4.72)	0.120 (0.30)	0.006 (1.19)	0.009 (0.34)	0.598 (1.46)	0.002 (0.54)	-0.058*** (-4.25)	-0.017 (-0.04)
ΔY		0.058 (1.09)	10.690*** (6.07)		0.134 (0.86)	9.164*** (3.90)		0.057 (0.95)	11.100*** (5.58)
ΔE	0.008** (2.94)		0.162 (0.40)	0.014** (3.08)		-0.690 (-1.71)	0.009* (2.45)		0.283 (0.61)
ΔT	0.044*** (5.28)	-0.002 (-0.07)		-0.024 (-0.76)	0.083 (0.46)		0.042*** (3.81)	0.006 (0.15)	
N	6898	6898	6898	1380	1380	1380	5520	5520	5520

Note: values in parentheses are t-values. Significance levels: *, P<0.05; **, P<0.01; ***, P<0.001. Lag and lead lengths both 1.
The "xtwest" Stata command was applied for the ECM-based panel cointegration test (Persyn and Westerlund, 2008)

The estimated long-run ECM coefficients are presented in Table 4.

Table 4: Estimated long-run ECM coefficients

Variable	FULL SAMPLE			OECD			Non -OECD		
	α_i^k	β_i^k	γ_i^k	α_i^k	β_i^k	γ_i^k	α_i^k	β_i^k	γ_i^k
Y	8.30	0.10	0.70	9.70	0.01	1.20	7.70	0.13	0.46
E	2.40	0.40	0.10	9.90	1.40	3.10	2.90	0.40	0.10
T	176.80	31.40	1.50	405.50	44.30	14.50	78.00	20.40	1.40

Note: It is difficult to test the significance of α_i^k , β_i^k , γ_i^k because the variances for these coefficients may not be available, so we did not estimate their standard errors. Y is the log of GDP.

According to our results, for the full sample and Non-OECD economies, a 1% increase in Y will increase E by 0.4 metric tons, which represents the long-term effect of Y on E over future periods. The increase of Y will cause deviations from the long-run equilibrium, causing E to be too high. E will then decrease to correct this disequilibrium, with the deviation decreasing by 13% (λ_i^E) in each subsequent time period for the full sample and 14% for Non-OECD economies. That is, E will decrease by an average of 0.4 metric tons in response, with the decrease occurring over successive future measurement intervals at a rate of 13% for the full sample and 14% for Non-OECD economies. A one-unit increase in T will increase E by 0.10 metric tons in both cases. To reestablish equilibrium E will then decrease by 0.10 metric tons over successive future measurement intervals at a rate of 13% for the full sample and 14% for Non-OECD economies per interval. For OECD countries, an increase of 1% in Y will increase E by 1.4 metric tons, while a one-unit increase in T will increase E by 3.1 metric tons. The return to equilibrium will occur at a rate of 5.5% per time interval.

The results of the short-run causality tests are presented in Table 5, where the direction of causal relationships is indicated by (→) for unidirectional causal relationships. According to our results, the relationship between Y and E exhibits bidirectional causality for OECD countries: i.e. a change in Y will affect E and a change in E will similarly affect Y. There is also a bi-directional relationship between T and Y for the full sample, implying that a change in Y will affect T and vice versa. For Non-OECD countries, E and Y are causally related to T, and there are unidirectional causal relationships from Y to T.

Table 5: Results of the short-run causality tests

Causality test	Null hypothesis	FULL SAMPLE	OECD	Non-OECD
$\Delta Y + \Delta T \rightarrow \Delta E$	$\phi^E_{i,j} = \delta^E_{i,j} = 0$	7.81**	26.54***	2.66
$\Delta Y \rightarrow \Delta E$	$\delta^E_{i,j} = 0$	0.142** (2.77)	0.671*** (4.78)	0.023 (0.60)
$\Delta T \rightarrow \Delta E$	$\phi^E_{i,j} = 0$	0.0273 (0.76)	-0.446** (-2.61)	0.064 (1.57)
$\Delta E + \Delta T \rightarrow \Delta Y$	$\theta^Y_{i,j} = \phi^Y_{i,j} = 0$	15.51***	42.01***	7.04
$\Delta E \rightarrow \Delta Y$	$\theta^Y_{i,j} = 0$	0.005 (1.80)	0.0237*** (4.76)	0.0037 (0.99)
$\Delta T \rightarrow \Delta Y$	$\phi^Y_{i,j} = 0$	0.028*** (3.51)	0.122*** (3.77)	0.027* (2.46)
$\Delta E + \Delta Y \rightarrow \Delta T$	$\theta^T_{i,j} = \delta^T_{i,j} = 0$	30.16***	0.55	12.49***
$\Delta E \rightarrow \Delta T$	$\theta^T_{i,j} = 0$	-0.0007 (-0.19)	0.0024 (0.57)	-0.00007 (-0.02)
$\Delta Y \rightarrow \Delta T$	$\delta^T_{i,j} = 0$	0.0934*** (5.49)	-0.0140 (-0.62)	0.048*** (3.53)

Note: values in parentheses are t-values. Significance levels: *, P<0.05; **, P<0.01; ***, P<0.001.
For the co-joint test, we used the Wald-test (χ^2)

The main findings can be summarized as follows. There is strong bi-directional short-run causality between CO_2 and GDP for OECD countries. This is consistent with expectations, since the OECD experienced a significant increase in CO_2 emissions that was especially pronounced in certain countries over the studied period. Furthermore, the higher the country's GDP (and income), the greater the amount of CO_2 that is likely to be released via production and/or consumption.

Dinda and Coondoo (2006) also found cointegrating relationships between CO_2 and GDP for Eastern and Western Europe, Central America, Africa, Japan and Oceania. In addition, they found evidence for their panel as a whole that strongly points to the existence of bi-directional causality.

Finally, there is bi-directional causality between international trade and GDP for the full sample and uni-directional causality between the same variables for OECD countries. For Non-OECD countries, there are no direct effects of GDP and trade on emissions. This implies that neither GDP growth nor international trade seems to have any significant short-run effect on CO_2 emissions for Non-OECD countries.

4. Conclusions

In this paper, we analyzed cointegration and short-run causal relationships between per capita CO_2 emissions, per capita GDP and international trade based on a cross-country panel data set covering 150 countries during the 1960–2008 period. Our estimates are based on the full sample of countries, as well as on two separate sub-samples, comprising OECD and Non-OECD countries, respectively.

Using the unit root test procedure, we found that all three series (the logarithm of the per capita GDP, per capita CO_2 emissions and trade measure) follow $I(1)$ processes. These findings were then used to apply ECM-based panel cointegration tests (Westerlund, 2007). The robust *P*-values obtained from the panel tests indicated that the null hypothesis of no cointegration should be rejected for both the full sample and the Non-OECD countries, while group-mean tests indicated that the null hypothesis of no cointegration can be rejected only for Non-OECD countries. This suggests that per capita CO_2 emissions, per capita GDP and the measure of international trade are cointegrated. Consequently, there are long-run equilibrium relationships among these three variables for both the full sample and the Non-OECD sample. Our results are consistent with previous findings; Dinda and Coondoo (2006) found a cointegration relationship between CO_2 emissions and GDP for 88 countries between 1960 and 1990, while Al-Mulali (2011) found a long-run relationship between CO_2 emissions and GDP for MENA countries[7].

The possible existence of causal relationships among per capita CO_2 emissions, per capita GDP and international trade has also been tested. The results suggest that there are short-run bi-directional causal relationships between per capita GDP and trade, together with a causal relationship between

[7] MENA countries refers to Middle East and North African countries.

CO_2 emissions plus GDP and trade, for the full sample. These findings suggest that economic policies should address growth, international trade and environmental pollution simultaneously.

Differences in the direction of causality have been detected between the two sub-samples considered. In the OECD sample, our results suggest there is bi-directional causality between per capita GDP and CO_2 emissions. This implies that policymakers should consider CO_2 emissions and economic growth simultaneously. Our results are partially consistent with those of Coondo and Dinda (2002), who found a unidirectional causal relationship between CO_2 emissions and GDP for developed country groups in North America and Western Europe and a unidirectional causal relationship from GDP to CO_2 emissions for country groups of Central and South America, Oceania and Japan.

For OECD countries, our results suggest that there are also causal relationships from GDP and international trade to per capita CO_2 emissions, and from per capita CO_2 emissions and international trade to per capita GDP. Conversely, for Non-OECD countries there are two uni-directional relationships, from per capita GDP to international trade and from per capita CO_2 emissions and per capita GDP to international trade. The absence of causal relationships between per capita CO_2 emissions and per capita GDP in Non-OECD countries implies that we do not have clear evidence that GDP affects CO_2 emissions. In contrast, previous studies (e.g. Coondo and Dinda, 2002) have identified a bi-directional relationship between these variables for Asian and African countries in the 1960–1990 period.

We would like to stress that a comprehensive analysis in this field would require a study of relationships among income, trade, emissions and energy, specifying the type of energy used, the structural composition of GDP and available technology, among other factors. However, the empirical framework employed in this study could be used to estimate the short- and long-run elasticities of CO_2 emissions in disaggregated sectors, in order to calibrate the developed models and generate scenarios describing how openness policies might motivate businesses to adopt environmentally friendly and efficient technologies to reduce emissions.

Appendices

Appendix A: Descriptive Statistics

Table A1: Descriptive statistics of variables

Full sample							
Variable	Unit	Variance	Mean	Std. Dev.	Min	Max	Obs.
CO_2 emissions	m.t.	overall	0.931	1.452	0.000	18.390	N = 7350
		between		1.297	0.008	7.807	n = 150
		within		0.660	-4.138	13.362	T = 49
Log(GDP)	$	overall	8.222	1.298	4.522	11.637	N = 7348
		between		1.238	5.353	10.972	n = 150
		within		0.403	5.246	10.963	T-bar = 49
Int. Trade	share	overall	0.716	0.520	-0.149	5.866	N = 7350
		between		0.446	0.020	3.096	n = 150
		within		0.270	-0.410	4.078	T = 49

OECD countries							
Variable	Unit		Mean	Std. Dev.	Min	Max	Obs.
CO_2 emissions	m.t.	overall	2.392	1.535	0.140	11.050	N = 1470
		between		1.445	0.577	7.807	n = 30
		within		0.579	-0.705	5.635	T = 49
Log(GDP)	$	overall	9.745	0.617	7.498	11.406	N = 1470
		between		0.489	8.669	10.458	n = 30
		within		0.386	8.374	11.025	T = 49
Int. Trade	share	overall	0.508	0.410	0.394	3.243	N = 1470
		between		0.355	0.154	2.098	n = 30
		within		0.214	-0.120	1.750	T = 49

Non-OECD countries							
Variable	Unit		Mean	Std. Dev.	Min	Max	Obs.
CO_2 emissions	m.t.	overall	0.566	1.175	0.000	18.390	N = 5880
		between		0.963	0.008	5.959	n = 120
		within		0.679	-4.503	12.997	T = 49
Log(GDP)	$	overall	7.840	1.136	1.852	11.637	N = 5880
		between		1.061	5.353	10.972	n = 120
		within		0.417	1.548	10.581	T = 49
Int. Trade	share	overall	0.7628	0.508	1.035	4.432	N = 5880
		between		0.437	2.003	3.096	n = 120
		within		0.262	-36.286	3.670	T = 49

Note: *Overall* refers to the whole dataset. The total variation (around grand mean $\bar{x} = 1/NT\sum_i \sum_t x_{it}$) can be broken down into *within* variation over time for each individual country (around individual mean $\bar{x}_i = 1/NT\sum_t x_{it}$) and *between* variation across countries (for $\bar{x}$ around $\bar{x}_i$). The corresponding breakdown for the variance is

Within variance: $s_W^2 = \frac{1}{NT-1}\sum_i \sum_t (x_{it} - \bar{x}_i)^2 = \frac{1}{NT-1}\sum_i \sum_t (x_{it} - \bar{x}_i + \bar{x})^2$;

Between variance: $s_B^2 = \frac{1}{N-1}\sum_i (x_i - \bar{x})^2$;

Overall variance: $s_O^2 = \frac{1}{NT-1}\sum_i \sum_t (x_{it} - \bar{x})^2$.

The second expression for s_w^2 is equivalent to the first, because adding a constant does not change the variance, and it is used at times because $x_{it} - \bar{x}_i + \bar{x}$ is centered on $\bar{x}$, providing a sense of scale, whereas $x_{it} - \bar{x}_i$ is centered on zero.

Appendix B: Westerlund's ECM based Panel Cointegration Test

Cointegration is tested according to the following specifications:

$$E_{it} = \mu_i^E + \tau_i^E t + \delta_i^E Y_{it} + \gamma_i^E T_{it} + u_{it}$$

$$Y_{it} = \alpha_i + \tau_i^Y t + \beta_i^Y E_{it} + \gamma_i^Y T_{it} + \varepsilon_{it}$$

$$T_{it} = \nu_i + \tau_i^T t + \delta_i^T Y_{it} + \beta_i^T E_{i,t} + e_{it}$$

Appendix B

Table B1: Results of Westerlund's ECM based Panel Cointegration Tests

Model	Test	No Constant noTrend									With Constant and Trend								
		Full sample			OECD			Non-OECD			Full sample			OECD			Non-OECD		
		Value of the test	z-value	p-value	Value of the test	z-value	p-value	Value of the test	z-value	p-value	Value of the test	z-value	p-value	Value of the test	z-value	p-value	Value of the test	z-value	p-value
Y	G_t	0.103	17.351	1.000	1.386	14.464	1.000	-0.218	12.165	1.000	-2.571	-0.622	0.267	-2.400	0.837	0.799	-2.627	-1.279	0.101
	G_a	-0.044	12.945	1.000	0.277	6.111	1.000	-0.124	11.418	1.000	-8.879	7.941	1.000	-7.341	4.701	1.000	-9.479	6.206	1.000
	P_t	0.653	9.128	1.000	8.344	10.141	1.000	-1.188	6.831	1.000	-25.231	3.089	0.999	-8.836	4.095	1.000	-111.265	-95.566	0.000
	P_a	0.020	6.294	1.000	0.317	3.147	0.999	-0.044	5.487	1.000	-8.538	3.532	1.000	-5.335	4.175	1.000	-37.829	-44.310	0.000
E	G_t	-1.233	1.738	0.959	-1.138	1.275	0.899	-1.255	1.324	0.907	-2.555	-0.393	0.347	-2.636	-0.702	0.241	-2.532	-0.053	0.479
	G_a	-4.054	3.965	1.000	-2.548	3.281	1.000	-4.427	2.798	0.997	-9.796	6.41	1.000	-6.935	5.004	1.000	-10.507	4.670	1.000
	P_t	-26.994	-11.678	0.000	-6.188	-0.795	0.213	-25.051	-11.128	0.000	-31.694	-4.076	0.000	-13.505	-1.081	0.140	-28.461	-3.771	0.000
	P_a	-6.093	-8.982	0.000	-2.314	0.207	0.582	-6.458	-8.850	0.000	-10.296	0.347	0.636	-8.312	1.762	0.961	-10.448	0.063	0.525
T	G_t	-1.186	2.289	0.989	-0.137	6.506	1.000	-1.245	1.433	0.924	-1.986	7.907	1.000	-1.866	4.315	1.000	-2.337	2.785	0.997
	G_a	-4.042	3.991	1.000	-0.613	5.219	1.000	-3.587	4.481	1.000	-5.385	13.778	1.000	-2.723	8.151	1.000	-7.851	9.659	1.000
	P_t	-20.061	-6.461	0.000	2.795	5.966	1.000	-17.568	-5.496	0.000	-22.468	6.152	1.000	-4.085	9.362	1.000	-24.850	3.511	1.000
	P_a	-4.713	-5.533	0.000	0.860	3.753	1.000	-4.265	-3.947	0.000	-7.332	5.717	1.000	-1.831	7.014	1.000	-8.699	3.241	0.999

Appendix C: List of countries

Afghanistan
Albania
Algeria
Angola
Antigua and Barbuda
Argentina
Australia*
Austria*
Bahamas
Bahrain
Bangladesh
Barbados
Belgium*
Belize
Benin
Bermuda
Bhutan
Bolivia
Botswana
Brazil
Brunei
Bulgaria
Burkina Faso
Burundi
Cambodia
Cameroon
Canada*
Cape Verde
Central African Repub.
Chad
Chile*
China
Colombia
Comoros
Congo, Dem. Rep. of
Congo, Republic of
Costa Rica
Cote d'Ivoire
Cuba
Cyprus
Denmark*
Djibouti
Dominica
Dominican Republic
Ecuador
Egypt
El Salvador
Equatorial Guinea
Ethiopia
Fiji
Finland*
France*
Gabon
Gambia
Germany*
Ghana
Greece*
Grenada
Guatemala
Guinea
Guinea Bissau
Guyana
Haiti
Honduras
Hong Kong
Hungary*
Iceland*
India
Indonesia
Iran
Iraq
Ireland*
Israel*
Italy*
Ivory Coast
Jamaica
Japan*
Jordan
Kenya
Kiribati
Korea, Republic of*
Laos
Lebanon
Liberia
Luxembourg*
Macao
Madagascar
Malawi
Malaysia
Maldives
Mali
Malta
Mauritania
Mauritius
Mexico*
Mongolia
Morocco
Mozambique
Nepal
Netherlands*
New Zealand*
Nicaragua
Niger
Nigeria
Norway*
Oman
Pakistan
Panama
Papua New Guinea
Paraguay
Peru
Philippines
Poland*
Portugal*
Romania
Rwanda
Samoa
Sao Tome and Principe
Senegal
Seychelles
Sierra Leone
Singapore
Solomon Islands
Somalia
South Africa
Spain*
Sri Lanka
St. Kitts-Nevis
St.Vincent & Grenadines
Sudan
Suriname
Swaziland
Sweden*
Switzerland*
Syria
Taiwan
Thailand
Togo
Tonga
Trinidad and Tobago
Tunisia
Turkey*
Uganda
United Kingdom*
United States*
Uruguay
Vanuatu
Venezuela
Vietnam
Zambia
Zimbabwe

*Note:** indicates OECD countries, the rest are Non-OECD countries

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III

Growth, migration and unemployment across Swedish municipalities

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Abstract

Fundamental questions in economics are why some regions are richer than others, why their economic growth rates vary, whether their growth tends to converge and the key factors that contribute to the variations. These questions have not yet been fully addressed, but changes in the local tax base are clearly influenced by the average income growth rate, net migration rate, and changes in unemployment rates. Thus, the main aim of this paper is to explore in depth the interactive effects of these factors (and local policy variables) in Swedish municipalities, by estimating a proposed three-equation system. Our main finding is that increases in local public expenditures and income taxes have negative effects on subsequent local income growth. In addition, our results support the conditional convergence hypothesis, i.e. that average income tends to grow more rapidly in relatively poor local jurisdictions than in initially "richer" jurisdictions, conditional on the other explanatory variables.

JEL classification: R11, R23, E24, O47

Keywords: Growth, net migration, unemployment, local policy, convergence

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1. Introduction

A fundamental question in economics is why some regions are richer than others. Since seminal work by Barro and Sala-i-Martin (1992, 1995), the empirical literature on regional growth has largely focused on the so-called convergence hypothesis predicted by neo-classical growth theory as presented by Solow (1956). That is, if all regions are equal in all other relevant aspects, relatively poor regions tend to catch up with initially richer regions in terms of average incomes or gross regional product, leading to an equalization of incomes across regions over time. Examples of studies testing the convergence hypothesis include Barro (1991), Barro and Sala-i-Martin (1991), Blanchard and Katz (1992), Borjas *et al.* (1992), Mankiw *at al.* (1992), and Sala-i-Martin (1996). Based on Swedish data, Persson (1997) and Aronsson *et al.* (2001) find evidence of convergence in per capita income across counties, and Lundberg (2003, 2006) finds evidence of convergence across municipalities. However, other studies find divergence, among others Romer (1986, 1990), Lucas (1988) and Scott (1989).

One of the difficulties of interpreting results from regressions of average income or gross regional product growth is that they may reflect changes in populations, the composition of the labor force and/or technological changes. Barro and Sala-i-Martin (1995) estimate an equation for average income growth with a systematic part depending on the average income, measured at the beginning of a given period, and the rate of net migration. To avoid endogeneity problems, due to potential interactions between these variables, the rate of net migration is instrumented. Other studies have also included the initial unemployment rate as an explanatory variable for regional growth. For instance, in an analysis of Swedish data, Lundberg (2003) finds that the unemployment rate, measured at the beginning of a given period, has a negative impact on the subsequent average income growth. Fagerberg *et al.* (1997) broaden the perspective by adopting a framework that takes into account the interdependence between income growth, migration and employment. They find support for the hypothesis that factors that impact GDP per capita growth also impact employment growth, and vice versa.

In this paper, we analyze the regional growth pattern in Sweden in a setting influenced by Fagerberg *et al.* (1997) and Aronsson *et al.* (2001). Fagerberg *et al.* propose and estimate a simultaneous equations model with GDP per capita growth, employment growth and migration as endogenous variables, using data for 64 European regions in the 1980s, while Aronsson *et al.* explore determinants of regional income growth and net migration in Sweden from 1970 to 1995. We extend previous investigations by examining interactions between average income growth, net migration and changes in unemployment, together with effects of factors that influence disparities in these variables, based on data for Swedish municipalities from 1990 to 2007. A key issue addressed is the effects of local policy variables, such as local income tax rates and local public expenditures, on the local growth pattern. Therefore, a three-equation system is estimated, where the local tax base growth is represented by three dependent variables: the growth rate of average income, net migration and the change in unemployment rate. These three dependent variables are determined using functions based on local policy variables, such as the initial local income tax rate, total local public expenditures per capita and initial shares of total local public expenditures on child care, primary- and secondary-education, care for the elderly, and social care. In addition, differences in the initial endowment of human capital, political representation and stability in the local parliament, and the initial demographic structure in each municipality are controlled for. This three-equation system is estimated using a fixed effects panel data approach with three stage least squares (3SLS) regression.

Although the existing literature on regional and local growth is quite extensive, this paper adds to it in several ways. First, in comparison to Fagerberg *et al.* (1997) we employ a richer set of explanatory variables and focus on effects of local policy variables on regional growth, rather than effects of industrial structure. Second, this paper extends the framework applied in previous analyses based on Swedish data – Swedish counties, Persson (1999), Aronsson *et al.* (2001); Swedish municipalities, Lundberg (2003, 2006), and Andersson *et al.* (2007) — by taking changes in unemployment into account. This is an important aspect of our paper as it takes a broader perspective than earlier studies on regional growth based on Swedish data, allowing us to relate estimates of average income growth to changes in both labor supply and unemployment rates. Third, our dataset covers a longer timeframe than previous studies based on Swedish municipalities, e.g. Lundberg (2003, 2006) and Andersson *et al.* (2007).

Regional disparities in local tax bases (and hence average incomes, migration and unemployment rates) have been on the Swedish political agenda for decades. One reason for this is that Swedish municipalities are the main providers of welfare services (such as child care, primary and secondary education, and care for the elderly), which are mainly financed by a proportional income tax and through a redistribution system. Thus, the local tax base affects the local governments' abilities to provide these services, which depend in the long-term on the growth of per capita income and the success of the municipality and local private sector in attracting labor (net immigration) and creating jobs (low unemployment). In this respect, Sweden is a particularly interesting case to study, as high-quality data are available and the country has a strongly decentralized public sector with autonomous local authorities.

The rest of the paper is structured as follows. In Section 2 we present stylized facts regarding the Swedish situation and describe changes in municipal tax bases, average income growth, net migration and unemployment rates from 1990 to 2007. Section 3 describes the applied empirical procedures and Section 4 the applied data. Results and interpretations are given in Section 5 and final conclusions and discussion are presented in Section 6.

2. Background and stylized facts regarding Sweden and the local growth pattern between 1990 and 2007

As mentioned in the introduction, in Sweden the local governments are the main providers of child care, primary and secondary schooling, care for the elderly and other social care. These services are primarily financed through a proportional local income tax, which the local authorities are free to adjust. This means that changes at the local level in average income growth and net migration, in combination with changes in employment rates, affect the local per capita tax base and, consequently, the local authority's ability to finance the public services they are obliged by the central government to provide. To equalize local per capita tax bases, the central government has tried in various ways to equalize economic opportunities across regions by implementing a redistribution system and providing targeted subsidies to the private sector. For instance, grant-in-aid provisions for the regional and local public sector have been introduced to compensate regions and municipalities with relatively small per capita tax bases, together with location and transportation subsidies to stimulate the local private sector. In addition, the national government has tried to improve the balance of local conditions through the strategic location of national institutions, such as large numbers of new university colleges, and re-location of government agencies. Although it seems reasonable to believe that the location of new universities has had a significant impact on individual municipalities, it has been difficult to find empirical evidence of any effect of this policy on average income growth, see Lundberg (2003).

Despite these efforts, as in many other countries, the regional disparities in local per capita tax bases and hence average income levels, income growth, net migration and unemployment rates have been, and remain, substantial.

The geographical distribution of relative local per capita tax bases in 2007 is shown in Figure 1. Figure 1 illustrates the geographical disparities in local per capita tax bases (in 2007) in relative values, ranging from 1 (for Årjäng, the municipality with the lowest relative tax base) to 2.12 (for Danderyd, the municipality with the highest relative tax base, 2.12-fold higher than that of Årjäng).

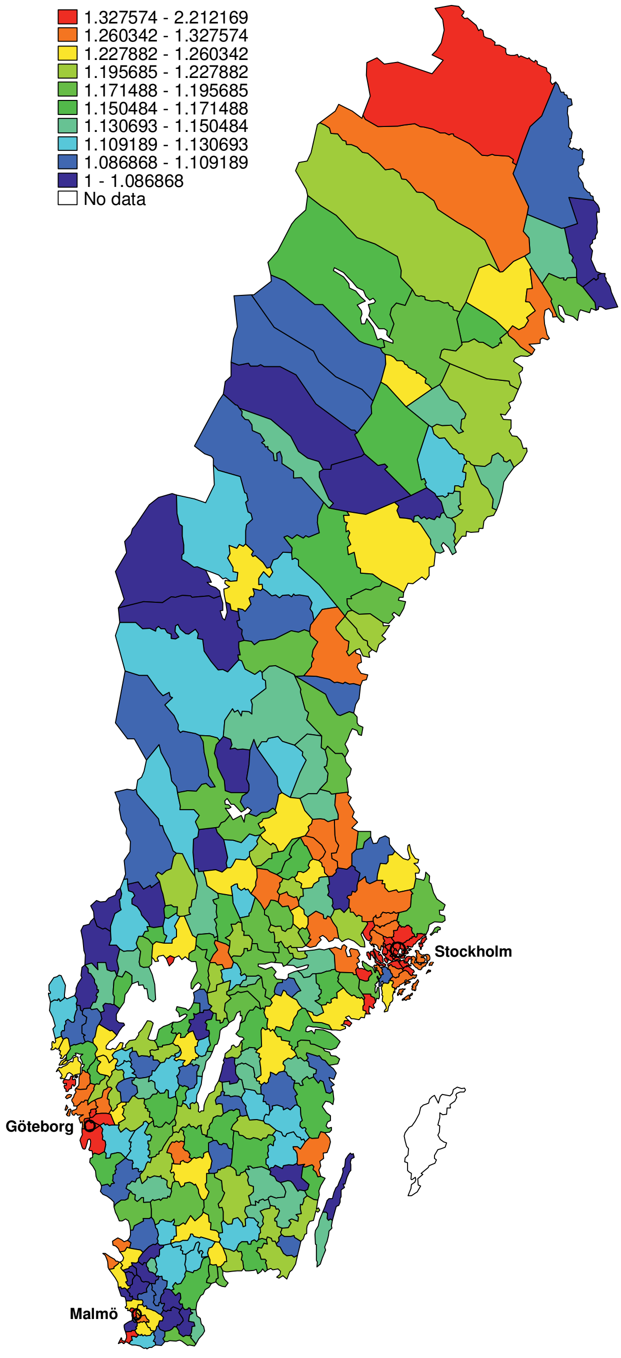

Figure 1: Relative per capita tax bases of Swedish municipalities, 2007

Municipalities with high per capita tax bases are clearly clustered in the highly urbanized areas of Stockholm, Göteborg and Malmö, as well as the most northerly county, Norrbotten, while municipalities with the smallest per capita tax bases are concentrated in western parts of Sweden.
It also shows that unemployment rates are highest and net immigration rates lowest in the already sparsely populated northern areas (such as Norrbotten, Västerbotten, and Värmland), while the situation is reversed in Stockholm and the surrounding area.

A key question is whether the disparities illustrated in Figure 1 have remained constant in recent decades, or local per capita tax bases have tended to converge (or diverge). To assess these possibilities, the correlation between the relative per capita tax bases of the municipalities in 1992 and 2007 is shown in Figure 2. The municipalities with the highest relative tax bases in 1992 (about 2.5 times higher than the lowest) still had the highest relative tax base in 2007 (2.2 times higher than the lowest). The large number of observations below and to the right of the 45 degree line suggests that the distribution in local per capita tax bases has become more compressed over time, i.e. the disparities in local per capita tax bases have decreased. Moreover, the correlation between initial per capita tax bases in 1992 and subsequent per capita tax base growth between 1992 and 2007 (Figure 3) suggests convergence over time, i.e. tax base growth since 1992 has been relatively low in municipalities with initially high relative per capita tax bases and vice versa. A simple OLS regression of initial per capita tax base in 1992 against local per capita tax base growth between 1992 and 2007 shows that this negative correlation is highly significant (*t*-value, -10.98) but very weak, with a parameter estimate of -0.011 (over 15 years), equating to an annual convergence rate of 0.07 percent.

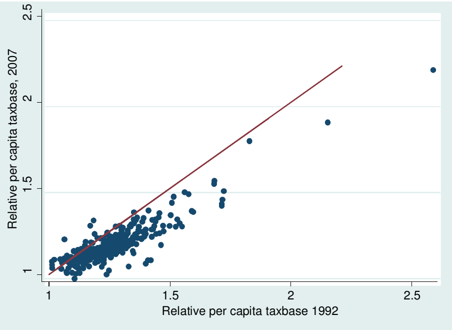

Figure 2: Correlation between the municipalities' relative per capita tax bases in 1992 and 2007

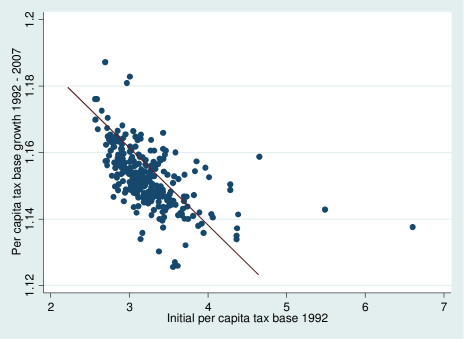

Figure 3: Correlation between the municipalities' relative per capita tax bases in 1992 (in logarithm) and subsequent growth in their per capita tax bases during 1992-2007

3. Empirical set up

Although the descriptive statistics presented and discussed above are interesting, they provide little indication of the processes responsible for the disparities. One way to broaden the analysis and acquire a better understanding of the determinants of local per capita tax base growth is to analyze effects of factors influencing specific components of the local tax base. Thus, here we examine potential determinants of average income growth, net migration and changes in unemployment rates. For instance, Aronsson *et al.* (2001) and Lundberg (2003) divided local tax base growth into two components, the average growth of income among the employed and net migration[1], thereby enabling exploration of correlations between parameter estimates of the average income growth equation and changes in both labor supply and the composition of the labor force. Fagerberg *et al.* (1997) took the decomposition a step further, suggesting that three components (average income growth, net migration and employment growth) are interdependent. They argued that relatively high average income growth is likely to lead to more jobs, thereby enhancing economic opportunities generally, and both net immigration and employment rates specifically. They assumed that migration influences income growth through its effects on labor supply and the composition of the labor force, expecting the relative productivity of immigrants (their endowments of human capital) to be positively related to the subsequent average income growth rate. Thus, Fagerberg *et al.* (1997) simultaneously estimated a three-equation system with average income growth, net migration and employment growth as endogenous variables.

Here, we estimate a system of three equations using the average income growth rate, net migration rate and changes in the unemployment rate as endogenous variables. By examining determinants of

[1] Net migration was used as a measure of population change.

these three variables simultaneously we take the analysis of Aronsson *el al.* (2001) and Lundberg (2003) a step further. We also apply a less structural approach than Fagerberg *et al.* (1997), while still taking the interdependence between average income growth, net migration rate and changes in the unemployment rate into account. That is, our goal is to recover key parameters of interest using exogenous within-sample variation with as few structural assumptions as possible. More formally, using $y_{i,t}$, $m_{i,t}$, $u_{i,t}$ and $x_{i,t-T}$ to denote the average income growth rate in municipality i between times $t-T$ and t, the net migration rate, the change in the unemployment rate, and a vector of relevant explanatory variables at time $t-T$, respectively, we simultaneously estimate the following system of equations:

$$y_{i,t} = f^{y}(x_{i,t-T}) \qquad (1a)$$

$$m_{i,t} = f^{m}(x_{i,t-T}) \qquad (1b)$$

$$u_{i,t} = f^{u}(x_{i,t-T}) \qquad (1c)$$

where $f^{y}(\cdot)$, $f^{m}(\cdot)$ and $f^{u}(\cdot)$ are linear form of the explanatory variables. Before presenting the full empirical model to be estimated, let us discuss factors that are potentially important determinants of local per capita tax base growth, i.e. the factors that should be included in $x_{i,t-T}$.

In an early paper Helms (1985) found that increases in local public consumption and redistribution negatively affect the subsequent growth rate while investments in roads, communication and human capital have positive effects. This suggests that different types of local public expenditures may have different effects. Other studies that have considered effects of local and national policy decisions include, among others, Glaeser *et al.* (1995) and Aronsson *et al.* (2001). To analyze effects of allocations of local government budgets between different locally provided public services on subsequent local growth patterns, the shares of local public expenditures on childcare, education, family care, care for the elderly, and culture and recreational services are included in the x vector. Local public expenditures on rescue services, business activities, and subsidies to political parties and the political process, are excluded and thus constitute the reference case. Based on Helms' results, it seems reasonable to expect, *a priori*, investments in human capital (expenditures on education) along with child and elderly care expenditures (which allow the working population to work longer hours) to have a positive effect on the subsequent average income growth and reduce unemployment. The other expenditure shares may have either positive or negative effects. In order to control for the size of the local public sector per capita, total local public expenditures and the initial local income tax rate[2] (measured as the sum of municipal and county rates) are included. In general, total public expenditures could have a positive effect on employment as labor is needed to provide public services. The tax rate may also reflect incentives to supply labor. However, public expenditures could also crowd out private investments, thereby reducing subsequent average income growth and employment rates, in accordance with Helms' finding that increases in taxes negatively affect growth.

Various proxies for "economic opportunities", such as the initial average income level, endowments of human capital and unemployment rates have been used as explanatory variables in growth regressions, for instance by Treyz *et al.* (1993), Westerlund and Wyzan (1995), Fagerberg *et al.* (1997), Aronsson *et al.* (2001), and Davies *et al.* (2001). A negative correlation between the initial income level and subsequent income growth rate is taken as evidence in favor of the conditional income convergence hypothesis. High average income levels and high endowments of human capital often signal social stability, making regions with these features attractive for migrants, and thus suggesting a positive correlation between these variables and subsequent net immigration rates. Moreover, the shares of the population with relatively high educational levels are expected to be positively correlated with the subsequent growth rate. However, previous studies have noted that apparent effects of educational levels may vary depending on the measures used, see for instance Di Liberto (2008), Barro and Lee (1994), Islam (1995), Pérez, *et al.* (1996), Pérez and Serrano (1998), Petrakis and Stamatakis (2002), and Pereira and Aubyn (2009). Therefore, following Lundberg (2003) we have included two measures of education: the initial share of the population with secondary education but less than three years of postsecondary education and the initial share of the population with more than three years of postsecondary education. The reference case is then the share of the population with secondary education or less. Employment (or unemployment) rates may be considered as indicators of the probability that a potential migrant would receive the average income level in a specific region. Hence, the initial unemployment rate is expected to be negatively correlated with both subsequent net immigration and average income growth rates.

Another potentially important determinant of economic growth we have included is the political stability in the local parliament (Barro, 1991; Glaeser *et al.,* 1995), which is expected to have a positive impact on the growth rate (Roubini and Sachs, 1989a, 1989b; Alesina and Perotti 1995; Alesina *et al.*, 1996). In addition, we control for the percentage of seats in the local parliament held by liberal and conservative parties and the municipalities' demographic structure. Finally, Westerlund and Wyzan (1995) found indications of differences in migration patterns between the major city

[2] Tax rates may also reflect incentives to supply labor.

areas and the rest of Sweden, implying important structural differences across regions. Therefore, a distinction is made between the major city areas (Stockholm, Göteborg and Malmö) and the rest of the country. The data set used, definitions of the variables, and full specifications of the empirical model to be estimated are presented in the next section.

4. Data source, variable definitions and empirical specification

The data used in this paper originate from Statistics Sweden (SCB) spanning from 1992 to 2007. During that period, the number of municipalities varied from 286 in 1992 to 290 in 2007. Gotland is excluded as the county and municipal levels coincide, making it difficult to separate county and municipal expenditures. In total, our estimates are based on 1,710 observations[3].

Starting with the three dependent variables, the average income growth rate ($y_{i,t}$) is defined as $\ln(Y_{i,t} / Y_{i,t-T})$, where $Y_{i,t}$ is the average income in year t of people aged 20 or above in municipality *i*. Here, $T = 10$ years. This seems a reasonable time interval in which to evaluate effects of the selected policy variables on local growth, migration rates and unemployment rates, which may only become fully apparent after substantial time (often many years). Thus, in previous empirical studies on regional growth the applied datasets have been divided into similar time intervals, for instance 5 years (Aronsson *et al.,* 2001), 9 years (Lundberg, 2005) or ≥ 10 years (Persson, 1997; Rey and Montori, 1999).

Net immigration between $t-T$ and t is defined as $m_{i,t} = \left(\sum_{l=t-T}^{l=t} mig_{i,l}\right) / \, pop_{i,t-T}$, where mig is the net number of immigrants and pop is the total population. Changes in the unemployment rate ($u_{i,t}$) are expressed as $\ln\left(Unemp_{i,t} / Unemp_{i,t-T}\right)$, where $Unemp$ is the number of unemployed divided by the population aged 25 to 64.

Initial endowments of human capital are captured by two variables: the share of the population aged 25 to 64 with less than 3 years post-secondary education ($Edulow_{i,t-T}$), and the share of the population aged 25 to 64 with more than 3 years post-postsecondary education ($Eduhigh_{i,t-T}$). Local policy variables are represented by the shares of local public expenditures allocated to: child care ($Exchil_{i,t-T}$); elementary, high school, and adult education ($Exedu_{i,t-T}$); family care, family

[3] We have information on 283 municipalities from 1990 to 1994, 287 municipalities from 1995 to 1998, 288 municipalities from 1999 to 2002, and 288 municipalities from 2003 to 2007. As we use data from 1990 and 1991 only as instruments, the total number of observations in our data is 2,276 while the second stage regression is based on 1,710 observations.

counseling, care and treatment of addiction, alcohol counseling etc. $(Exfam_{i,t-T})$; elderly care $(Exeld_{i,t-T})$; cultural and recreational services $(Excul_{i,t-T})$; per capita total local public expenditures $(Exp_{i,t-T})$; and the local income tax rate $(Tax_{i,t-T})$. As mentioned above, local public expenditures on rescue services, business activities, and subsidies to political parties and the political process are excluded and therefore constitute the reference case.

Political fragmentation in the local council is captured by a Herfindahl index $(Herf_{i,t-T})$, defined as $\sum_{p=k}^{P} SH_p^2$, where SH_p is the share of representatives from party p in the local parliament. Political ideology is measured by the share of the seats in the local parliament held by representatives of the liberal and conservative parties, defined here as the Conservative party, Liberal party, Christian Democratic party, and the Center party $(Cons_{i,t-T})$. Demographic characteristics are controlled for by including the shares of the population aged 0 to 6 $(Age06_{i,t-T})$, 7 to 15 $(Age715_{i,t-T})$, more than 65 $(Age65_{i,t-T})$, and the population density $(Dens_{i,t-T})$. All monetary variables are deflated by consumer prices index using 2005 as base year.

Table A1 in the Appendix provides more detailed definitions, Table A2 presents descriptive statistics and Table A3 displays a correlation matrix of all variables for 1992-2007, which clearly shows that the municipalities' expenditure shares are correlated with the population shares. The correlations are especially high between $Age06_{i,t-T}$ and $(Age65_{i,t-T})$, as well as between the expenditure shares $Exchil_{i,t-T}$ and $Exeld_{i,t-T}$. This might be problematic from an econometric perspective, because if all these variables were included in the same regression it would be difficult to accurately estimate the parameters. We will return to this issue when we discuss the results in Section 5.

Based on the discussion in Section 3, and the above definitions of the variables, the average income growth, net migration, and changes in unemployment rates are assumed to develop according to[4]:

[4] We use a log model because it is not affected by the scale of the independent variables when estimating relative changes in dependent variables. Taking logs also reduces extrema in the data and effects of outliers.

$$
\begin{aligned}
&y_{i,t} = \\
&(\alpha_i^y + \alpha_d^y D) + (\beta^y + \beta_d^y D)\ln Y_{i,t-T} + (\gamma_{un}^y + \gamma_{un,d}^y D)\ln Unemp_{i,t-T} + (\gamma_{tax}^y + \gamma_{tax,d}^y D)\ln Tax_{i,t-T} + \\
&(\gamma_{\exp}^y + \gamma_{\exp l,d}^y D)\ln Exp_{i,t-T} + (\gamma_{cul}^y + \gamma_{cul,d}^y D)\ln Excul_{i,t-T} + (\gamma_{edu}^y + \gamma_{edu,d}^y D)\ln Exedu_{i,t-T} + \\
&(\gamma_{chil}^y + \gamma_{chil,d}^y D)\ln Exchil_{i,t-T} + (\gamma_{fam}^y + \gamma_{faml,d}^y D)\ln Exfam_{i,t-T} + (\gamma_{eld}^y + \gamma_{eld,d}^y D)\ln Exeld_{i,t-T} + \\
&(\gamma_{low}^y + \gamma_{low,d}^y D)\ln Edulow_{i,t-T} + (\gamma_{high}^y + \gamma_{high,d}^y D)\ln Eduhigh_{i,t-T} + (\gamma_{cons}^y + \gamma_{cons,d}^y D)\ln Cons_{i,t-T} + \\
&(\gamma_{herf}^y + \gamma_{herf,d}^y D)\ln Herf_{i,t-T} + (\gamma_{a06}^y + \gamma_{a06,d}^y D)\ln Age06_{i,t-T} + (\gamma_{a715}^y + \gamma_{a715,d}^y D)\ln Age715_{i,t-T} + \\
&(\gamma_{a65}^y + \gamma_{a65,d}^y D)\ln Age65_{i,t-T} + (\gamma_{dens}^y + \gamma_{dens,d}^y D)\ln Dens_{i,t-T} + \varepsilon_{i,t}^y
\end{aligned} \tag{2a}
$$

$$
\begin{aligned}
&m_{i,t} = \\
&(\alpha_i^m + \alpha_d^m D) + (\beta^m + \beta_d^m D)\ln Y_{i,t-T} + (\gamma_{un}^m + \gamma_{un,d}^m D)\ln Unemp_{i,t-T} + (\gamma_{tax}^m + \gamma_{tax,d}^m D)\ln Tax_{i,t-T} + \\
&(\gamma_{\exp}^m + \gamma_{\exp l,d}^m D)\ln Exp_{i,t-T} + (\gamma_{cul}^m + \gamma_{cul,d}^m D)\ln Excul_{i,t-T} + (\gamma_{edu}^m + \gamma_{edu,d}^m D)\ln Exedu_{i,t-T} + \\
&(\gamma_{chil}^m + \gamma_{chil,d}^m D)\ln Exchil_{i,t-T} + (\gamma_{fam}^m + \gamma_{faml,d}^m D)\ln Exfam_{i,t-T} + (\gamma_{eld}^m + \gamma_{eld,d}^m D)\ln Exeld_{i,t-T} + \\
&(\gamma_{low}^m + \gamma_{low,d}^m D)\ln Edulow_{i,t-T} + (\gamma_{high}^m + \gamma_{high,d}^m D)\ln Eduhigh_{i,t-T} + (\gamma_{cons}^m + \gamma_{cons,d}^m D)\ln Cons_{i,t-T} + \\
&(\gamma_{herf}^m + \gamma_{herf,d}^m D)\ln Herf_{i,t-T} + (\gamma_{a06}^m + \gamma_{a06,d}^m D)\ln Age06_{i,t-T} + (\gamma_{a715}^m + \gamma_{a715,d}^m D)\ln Age715_{i,t-T} + \\
&(\gamma_{a65}^m + \gamma_{a65,d}^m D)\ln Age65_{i,t-T} + (\gamma_{dens}^m + \gamma_{dens,d}^m D)\ln Dens_{i,t-T} + \varepsilon_{i,t}^m
\end{aligned} \tag{2b}
$$

$$
\begin{aligned}
&u_{i,t} = \\
&(\alpha_i^u + \alpha_d^u D) + (\beta^u + \beta_d^u D)\ln Y_{i,t-T} + (\gamma_{un}^u + \gamma_{un,d}^u D)\ln Unemp_{i,t-T}(\gamma_{tax}^u + \gamma_{tax,d}^u D)\ln Tax_{i,t-T} + \\
&(\gamma_{\exp}^u + \gamma_{\exp l,d}^u D)\ln Exp_{i,t-T} + (\gamma_{cul}^u + \gamma_{cul,d}^u D)\ln Excul_{i,t-T} + (\gamma_{edu}^u + \gamma_{edu,d}^u D)\ln Exedu_{i,t-T} + \\
&(\gamma_{chil}^u + \gamma_{chil,d}^u D)\ln Exchil_{i,t-T} + (\gamma_{fam}^u + \gamma_{faml,d}^u D)\ln Exfam_{i,t-T} + (\gamma_{eld}^u + \gamma_{eld,d}^u D)\ln Exeld_{i,t-T} + \\
&(\gamma_{low}^u + \gamma_{low,d}^u D)\ln Edulow_{i,t-T} + (\gamma_{high}^u + \gamma_{high,d}^u D)\ln Eduhigh_{i,t-T} + (\gamma_{cons}^u + \gamma_{cons,d}^u D)\ln Cons_{i,t-T} + \\
&(\gamma_{herf}^u + \gamma_{herf,d}^u D)\ln Herf_{i,t-T} + (\gamma_{a06}^u + \gamma_{a06,d}^u D)\ln Age06_{i,t-T} + (\gamma_{a715}^u + \gamma_{a715,d}^y D)\ln Age715_{i,t-T} + \\
&(\gamma_{a65}^u + \gamma_{a65,d}^u D)\ln Age65_{i,t-T} + (\gamma_{dens}^u + \gamma_{dens,d}^u D)\ln Dens_{i,t-T} + \varepsilon_{i,t}^u
\end{aligned} \tag{2c}
$$

Here, the α's, β's and γ' s are parameters to be estimated, and the ε 's are error terms, in ten-year windows with $T = 10$ years and $t = 2004,\ 2005,\ 2006$ and 2007. D is a dummy variable taking the value 1 for the major city areas, otherwise zero. This specification gives different parameter estimates for the major city areas compared with the rest of the country, in accordance with previous findings using Swedish data, see Westerlund and Wyzan (1995).

This set up allows various hypotheses to be easily tested. For instance, $\beta^y < 0$ indicates that the conditional convergence hypothesis is valid for municipalities outside of the major city areas Stockholm, Göteborg and Malmö, while $\beta^y + \beta_d^y < 0$ is consistent with conditional convergence in the major city areas.

Equations (2a), (2b) and (2c) are estimated by applying an unbalanced fixed effects panel data approach[5] using three-stage least squares (3SLS) regression. As model (2a) includes the initial income per capita as an explanatory variable, it cannot be estimated consistently with OLS, due to endogeneity problems associated with correlations between the lagged dependent variable and the fixed effects. Therefore, the three equations are simultaneously estimated via 3SLS. The instrumentation of $Y_{i,t-T}$ is influenced by Arellano and Bond (1991) and more generally by Baltagi (1995), which essentially involves using lagged variables as instruments, with efficiency gained by expanding the set of instruments over time. That is, $Y_{i,t-2}$ and $Y_{i,t-3}$ are used as instruments for $Y_{i,t}$; $Y_{i,t-1}$, $Y_{i,t-2}$ and $Y_{i,t-3}$ as instruments for $Y_{i,t+1}$ and so on. This procedure exploits the validity of using $Y_{i,t-2}$, $Y_{i,t-3},...,$ as instruments for the lagged dependent variable $Y_{i,t}$ that generate consistent and efficient estimates of the parameters of interest.

5. Results

Parameter estimates of the three equations (2a), (2b) and (2c) are presented in Table 1. The "Basic" and "Major city" columns respectively show estimated parameters of the indicated variables for municipalities located outside the major city areas (Stockholm, Göteborg and Malmö), and the major city areas (the latter calculated as the sum of the "Basic" and the "Dummy" estimates of the respective variables). In the following text, the data presented in parentheses are the parameters obtained and corresponding *t*-values.

Let us start by discussing conditional convergence, which should be reflected (if present) in a negative parameter for the relationship between the initial average income level (Y) and subsequent average income growth (y). According to the estimates in Table 1, the initial average income is significantly, negatively related to the subsequent average income growth (-0.570, *t*-value -5.34) for municipalities outside the major city areas. That is, among these municipalities, those with relatively low initial income levels tend to grow faster, *ceteris paribus*, than municipalities with relatively high initial income levels, conditional on the other explanatory variables in the model. This result also holds for the major city areas (the sum of the "Basic" and "Dummy" parameter estimates is -0.483, *t*-value -7.13). These results are consistent with findings of previous studies based on both Swedish data (e.g. Persson, 1997; Aronsson *et al.*, 2001; Lundberg, 2003), and data on U.S. states (e.g. Barro and Sala-i-Martin, 1992, 1995). Our results suggest a convergence rate of about 5% per year, faster

[5] This approach allows us to control for all unit-specific factors whether observable or unobservable that are constant over time. Including lagged dependent variables in a model can also control to a large extent for many omitted variables.

than rates frequently reported in the growth literature. One reason for this could be the inclusion of fixed effects in addition to other covariates, which also implies that each municipality is converging at a specific rate, rather than the average rate within the country.

Table 1: Estimated parameters for relationships between the indicated variables

	BASIC	MAJOR CITY	BASIC	MAJOR CITY	BASIC	MAJOR CITY
Dependent variable	$y_{i,t}$		$m_{i,t}$		$u_{i,t}$	
$Y_{i,t-T}$	-0.570 (-5.34)	-0.483 (-7.13)	0.287 (2.99)	0.172 (2.83)	0.457 (1.09)	0.154 (0.58)
$Unemp_{i,t-T}$	0.065 (5.90)	0.044 (6.49)	-0.009 (-3.09)	-0.004 (-0.75)	0.376 (8.67)	0.432 (16.06)
$Tax_{i,t-T}$	-0.053 (-0.50)	0.006 (0.09)	-0.240 (-2.52)	-0.153 (-2.53)	0.050 (0.12)	0.062 (0.23)
$Exp_{i,t-T}$	-0.052 (-1.80)	-0.030 (-1.74)	-0.052 (-2.02)	-0.030 (-1.97)	0.112 (1.00)	0.60 (0.89)
$Excul_{i,t-T}$	0.003 (0.29)	0.000 (0.05)	-0.006 (-0.57)	-0.002 (-0.27)	0.001 (0.01)	-0.004 (-0.13)
$Exedu_{i,t-T}$	-0.048 (-2.20)	-0.027 (-2.09)	-0.033 (-1.66)	-0.018 (-1.55)	0.096 (1.11)	0.052 (1.01)
$Exchil_{i,t-T}$	-0.057 (-3.31)	-0.033 (-3.21)	-0.011 (-0.72)	-0.005 (-0.52)	0.030 (0.45)	0.025 (0.61)
$Exfam_{i,t-T}$	-0.006 (-0.91)	-0.004 (-0.99)	-0.010 (-1.64)	-0.007 (-1.87)	-0.001 (-0.04)	0.001 (0.08)
$Exeld_{i,t-T}$	0.019 (1.34)	0.006 (0.68)	-0.005 (-0.36)	-0.003 (-0.36)	-0.064 (-1.13)	-0.046 (-1.29)
$Edulow_{i,t-T}$	-0.046 (-1.36)	-0.026 (-1.28)	-0.082 (-2.71)	-0.074 (-4.01)	0.084 (0.64)	0.068 (0.85)
$Eduhigh_{i,t-T}$	0.131 (2.34)	0.067 (2.05)	0.075 (1.50)	0.029 (1.00)	-0.776 (-3.54)	-0.467 (-3.63)
$Cons_{i,t-T}$	0.001 (0.08)	-0.007 (-0.76)	0.046 (3.13)	0.023 (2.69)	0.000 (0.00)	0.026 (0.68)
$Herf_{i,t-T}$	0.013 (0.64)	0.011 (0.83)	-0.028 (-1.48)	-0.013 (-1.10)	0.051 (0.62)	0.028 (0.55)
$Age06_{i,t-T}$	-0.259 (-4.87)	-0.166 (-5.18)	0.101 (2.11)	0.069 (2.38)	0.023 (0.11)	0.027 (0.22)
$Age715_{i,t-T}$	-0.060 (-0.91)	-0.042 (-1.07)	-0.014 (-0.23)	-0.011 (-0.32)	0.324 (1.25)	0.152 (0.98)
$Age65_{i,t-T}$	0.168 (2.12)	0.102 (2.05)	-0.014 (-0.20)	0.027 (0.61)	-0.470 (-1.51)	-0.235 (-1.20)
$Dens_{i,t-T}$	0.269 (3.18)	0.219 (4.26)	0.272 (3.57)	0.081 (1.75)	0.041 (0.12)	-0.040 (-0.20)
# of observations	1710		1710		1710	

Note: t-values in parentheses.

For further interpretation, it is useful to examine effects of the initial average income level (Y) on subsequent net migration rates (m) and changes in unemployment rates (u). Our results suggest that the initial average income level is positively related to the subsequent net migration rate for municipalities located both outside and in the major city areas (0.287, *t*-value 2.99 and 0.172, *t*-value 2.83, respectively), indicating that high average income levels make municipalities more attractive to migrants. If individuals generally migrate to areas with high average income levels (areas with high per capita productivity), the labor supply in those areas will increase, thereby negatively affecting the subsequent average income growth rate, provided there are no significant differences in other determinants. Hence, this suggests that labor mobility may contribute to the equalization of average income levels across municipalities as the labor force tends to migrate to municipalities with high average incomes. The speed of convergence (or equalization) could also be affected if the labor mobility affects the composition of the labor force. For instance, if those who migrate to relatively high income municipalities are less productive (e.g. have lower human capital)[6] than those who stay, the per capita productivity will decrease in municipalities with initially high average incomes, but increase in municipalities with initially low average incomes. This could partly explain the relatively high convergence rate we find across Swedish municipalities. Our model does not predict any significant relationship between the initial average income and subsequent changes in unemployment rates. Two other factors that may contribute to convergence, proposed by Aronsson *et al.*, relate to high capital mobility and the Swedish system for setting wages, which may make municipalities more homogeneous over time and compress the wage distribution, respectively.

The initial unemployment rate could be seen as an indicator of economic opportunities and future earning possibilities. The parameter estimates presented in Table 1 suggest that the initial unemployment rate $(Unemp)$ is significantly positively related to the subsequent average income growth (y) for municipalities both outside and in the major city areas (0.065, *t*-value 5.90 and 0.044, *t*-value 6.49, respectively). This suggests that average income growth rates tend to be higher in municipalities with high initial unemployment rates than in other municipalities. For further interpretation the relationships between the initial unemployment rate $(Unemp)$ and the subsequent net migration rate (m) and changes in unemployment rates (u) provide useful indications. Our model predicts that the initial unemployment rate $(Unemp)$ is negatively related to the subsequent net migration rate (m), in accordance with findings by Aronsson *et al.* (2001). The estimated parameter for this negative effect is only significant for municipalities outside the major city areas.

[6] Here, human capital includes both work experience and formal education. In Sweden, as in many other countries, the propensity to migrate across municipal borders decreases with age. Hence, among those who migrate individuals with low work experience are likely to be overrepresented.

However, we can take the analysis one step further than Aronsson *et al.* by incorporating the relationship between $(Unemp)$ and (u). The results presented in Table 1 suggest that these variables are highly significantly and positively correlated (0.376, *t*-value 8.67 and 0.432, *t*-value 16.06, respectively), indicating that growth in unemployment tends to be higher in municipalities with high initial unemployment rates than in other municipalities. One potential interpretation of the negative relationship between $(Unemp)$ and (m) (at least outside the major city areas), in combination with the positive correlation between $(Unemp)$ and (u), is that most individuals who tend to migrate from municipalities with initially high unemployment rates are employed, leading to increased unemployment rates within these areas. At the same time there seems to be a mismatch in the local labor markets. That is, as employed individuals tend to emigrate from areas with initially high unemployment rates the prospects for previously unemployed individuals finding employment should increase. However, if the skills of the unemployed do not match those required in the local labor market the emigration of employed individuals will reduce the labor supply and, consequently, raise incomes of employed individuals who have remained, as indicated by the positive correlation between $(Unemp)$ and (y). This interpretation is consistent with the positive (and highly significant) correlation between the initial unemployment rate and subsequent changes in unemployment rates. Hence, the work opportunities of those that remain deteriorate.

Turning to local policy variables, the local income tax rate is a factor that might influence migration between municipalities located in densely populated areas near the major cities, where a decision to move does not necessarily mean that the individual changes his/her place of work or social network. The local income tax rate may also crowd out private investments and thus have a negative effect on the subsequent average income growth. Our results do not provide any evidence suggesting that the initial income tax rate affects the subsequent average income growth significantly. However, our results suggest a negative relationship between the initial local income tax rate (Tax) and the subsequent net migration rate (m) for both municipalities outside the major city areas (-0.240, *t*-value -2.52) and municipalities in the major city areas (-0.153, *t*-value -2.53). No significant correlation is found between (Tax) and (u). Thus, we conclude the initial income tax rate is negatively related to the subsequent net migration rate but has no significant effects on either (y) or (u). A possible explanation is that the composition of the labor force is unaffected by the net emigration caused by high income tax rates.

From a theoretical perspective, local public expenditures could have either positive or negative effects on the subsequent growth rate. Results presented by Helms (1985) suggest that public investments in roads and education enhance growth while local public consumption crowds out private initiatives and hence negatively affects growth. Our model indicates that initial local public expenditures per capita (Exp) are negatively related to subsequent net migration rates (m) for both the major city areas (-0.030, *t*-value -1.97) and the rest of the country (-0.052, *t*-value -2.02), but has no significant effect on either the subsequent average income growth (y) or changes in unemployment rates (u). In terms of Helms' interpretations, this suggests that local public expenditures in Sweden are a mixture of investments in human capital and local public consumption, leaving both the subsequent average income growth rate and changes in unemployment unaffected. If increased public expenditure reduces average income growth, for which we have weak evidence, this might be due to the crowding-out effect that Helms also refers to (or other disincentives during revenue collection). In turn, lower income growth may signal worse employment prospects in the future, leading to emigration. This result is consistent with the findings of Fagerberg *et al.* (1996). Our results also indicate that although high initial local public expenditures tend to lead to net emigration, they may have little effect on the composition of the labor force.

It is of interest to consider whether, and if so to what extent, expenditure allocations by the local government affect subsequent growth patterns. Our results indicate that the shares of local public expenditure on cultural activities $(Excul)$, family care $(Exfam)$ and care for the elderly $(Exeld)$ have no significant effects on the subsequent growth pattern. However, shares of local public expenditures on both education $(Exedu)$ and child care $(Exchil)$ appear to be negatively related to subsequent average income growth (y). These results are surprising, as expenditures on both child care and education[7] could be viewed as investments in future productivity and thus are expected to enhance subsequent average income growth. A potential explanation is that these expenditures may also reflect social disparities between municipalities. In support of this hypothesis, the local public sector was the main provider of child care and both primary and secondary education during the study period, and although local authorities enjoy considerable autonomy from the central government, child care and education were (and are) highly regulated by national authorities. Thus, municipalities with "unfavorable" socio-economic backgrounds may have higher child care and school expenditures than other municipalities. That is, in order to achieve the standards set by national authorities, local public expenditures on child care and education might reflect socio-economic differences between municipalities. In addition, the findings could be due to a potential labor supply effect of children, i.e., parents of small and school-aged children cannot work as long hours as adults with no children. Moreover, as correctly pointed out by Aronsson *et al.* (2001), it is

[7] Remember that expenditures on education are measured as elementary, high school and adult education except university education.

generally difficult to interpret such results as local public expenditures and local income tax rates not only reflect current service levels, but may also give signals of future local policies. In fact, neither the counties nor the municipalities were required to balance their budgets every year during the study period, making it difficult to form expectations in advance about their effects on the subsequent local growth pattern[8]. Considering the impact of $(Exchil)$ on the net migration rate (m) and unemployment rate (u) gives no further guidance for the interpretation of these results. Nor do our results indicate any significant correlations between expenditures on elderly care $(Exeld)$ and average rates of growth (y), migration (m) or unemployment (u). These rather contradictory results may be explained by the lack of the ability, using our data, to evaluate how the quality of care and education provisions for children $(Exchil$ and/or $Exedu)$ in particular municipalities influence decisions of potential emigrants to remain in them. Given this inability, and lacking other plausible hypotheses, we simply conclude that our results indicate that local public expenditures and local income tax rates are negatively related to the local growth rate.

As mentioned in section 4 and displayed in Table A3, $Age06_{i,t-T}$ and $Age65_{i,t-T}$ are highly correlated with the expenditure shares $Exchil_{i,t-T}$ and $Exeld_{i,t-T}$, respectively. Although this is not surprising, it might be problematic from an econometric perspective as it makes it more difficult to accurately estimate parameters for these variables. We have tested the relationships using models with different specifications, excluding all expenditure shares, some of the expenditure shares, and the age distribution, etc. However, none of the variations significantly affected the signs of the parameter estimates or their significance level (data not shown).

The endowments and formation of human capital are often seen as major determinants of economic growth and wealth. Human capital is a complex concept and includes several unobservable components, but in empirical growth models it is often measured as the share of the population with a certain type of formal education, such as a university degree. For instance, results presented by Di Liberto (2008), Barro and Lee (1994), Islam (1995), Pérez *et al.* (1996), Pérez and Serrano (1998), Petrakis and Stamatakis (2002), and Pereira and Aubyn (2009) suggest that different levels of formal education have different effects on the subsequent growth pattern. Here, we use two measures: the share of the population aged 25 to 64 with less than 3 years post-postsecondary education $(Edulow)$, and the share of this population with more than 3 years post-postsecondary education $(Eduhigh)$. Although Sweden could be regarded as a highly educated country, in comparison to many other countries, there are large regional disparities in levels of formal

[8] Balanced budget requirements were imposed in 2000.

education. Shares of the population with at least three years postsecondary education are highest in or close to the major city areas and lowest in rural areas. The results presented in Table 1 suggest a negative relationship between $(Edulow)$ and (m), but no significant correlations between $(Edulow)$ and either (y) or (u). On the other hand, our model indicates a significantly positive relationship between $(Eduhigh)$ and the subsequent average income growth rate (y), and a significantly negative relationship between $(Eduhigh)$ and subsequent changes in the unemployment rate (u). These findings indicate that different levels of education have different impacts on the subsequent growth rate. Higher education appears to have positive effects[9], as the share of the population with at least three years of university education is positively related to the subsequent growth pattern. In contrast, the share of the population with only a secondary education is negatively related to the subsequent net migration rate. The latter result conflicts with findings presented by Aronsson *et al.* (2001) and Lundberg (2006), that human capital has no effect on subsequent net migration. To explore these relationships in more detail it would be interesting to differentiate between types of education, but we do not have access to pertinent information.

Other potentially important determinants of economic growth we tested are the political composition and stability of the local council. The results suggest that our indicator of political representation — the share of seats in the local parliament held by members of the Conservative Party, Centre Party, Liberals and Christian Democrats $(Cons)$ — is positively related to subsequent immigration (m) both inside and outside the major city areas (0.046, *t*-value 3.13 and 0.023, *t*-value 2.69, respectively), but has no significant effects on either (y) or (u). Given that conservatives and social democrats often collaborate at the local government level, the political agenda is in many respects different at the local level compared to the national level. In addition, our controls of effects of factors such as the initial average income level, endowments of human capital, income tax rates and local public expenditures exclude explanations for this result based on them. One tempting interpretation is that conservatives and liberals generally favor private initiatives and private entrepreneurs more strongly than their political opponents, but we do not have empirical support to confirm such an interpretation. In contrast to political representation, our model indicates that political stability, measured using a Herfindahl-index $(Herf)$, has no significant effect on the subsequent growth rate. This conflicts with findings of several previous studies.

[9] We have further examined this relationship by using different definitions of human capital and higher education. When we combine *Eduhigh* and *Edulow* into a single variable we find no significant effect of formal education on subsequent growth.

Finally, according to our estimates the measures of socio-economic and demographic structure mainly affect the subsequent average income growth and net migration rate. The initial share of inhabitants aged 0-6 $(Age06)$ is negatively correlated with (y) and positively correlated with the net migration rate (m) for municipalities situated both outside and inside the major city area. One potential explanation for these results is that some municipalities are more attractive than others to families with small children, and as the children grow up the attracted families tend to stay, thereby increasing the labor supply and (hence) negatively affecting (y). We also obtain a positive correlation between the initial share of inhabitants aged 65 and above $(Age65)$ and (y). This effect is more difficult to explain, we simply note that it is consistent with findings reported by Lundberg (2006). Population density $(Dens)$ also appears to have a significant, positive effect on the subsequent growth pattern in municipalities both outside and inside major city areas (3.18, *t*-value 0.269 and 0.219, *t*-value 4.26, respectively), as found by Aronsson *et al.* (2001). Population density is also positively associated with net migration for municipalities outside major city areas (0.272, *t*-value 3.57). However, there is no evidence that the population density $(Dens)$ affects the unemployment rate (u).

6. Conclusions

This paper explores effects of key determinants (the average income growth rate, net migration rate, and changes in unemployment rates) of changes in local tax bases of Swedish municipalities in an attempt to identify factors that may explain disparities among them in this respect. In recognition of the importance of these three variables, a three-equation system is estimated. Based on a dataset covering Swedish municipalities from 1992 to 2007, our results show that initial local public expenditures and income taxes are negatively related to the subsequent local growth pattern. Our results also support the conditional convergence hypothesis for average income levels across Swedish municipalities, i.e. that average income tends to grow more rapidly in initially "poor" local jurisdictions than in initially "richer" jurisdictions, conditional on the other explanatory variables

Moreover, we find that the share of individuals with more than three years formal postsecondary education is positively related to the subsequent net migration rate and negatively related to subsequent changes in unemployment rates, while the share of individuals with secondary education but less than three years postsecondary education is negatively related to the subsequent net migration rate.

Besides the findings that local public expenditures and taxes seem to have negative effects on subsequent local growth, one of the most interesting results of our study is a positive correlation between the initial unemployment rate and the subsequent average income growth. As the initial unemployment rate is positively correlated with the subsequent change in unemployment rate and negatively correlated with the subsequent net migration rate, we attribute this result to mismatches in the local labor market. This interpretation is enabled by the three-equation approach used in this paper.

Appendices

Table A1: Definitions

Variable	Description
Economic factors	
$Y_{i,t}$	Average income (thousand SEK) of employed individuals aged ≥20 years.
$Unemp_{i,t}$	Number of unemployed divided by the population aged 25 to 64.
$y_{i,t} = \ln(Y_{i,t} / Y_{i,t-T})$	Average income growth.
$m_{i,t} = \left(\sum_{l=t-T}^{l=t} mig_{i,l}\right) / pop_{i,t-T},$	Net migration rate.
$u_{i,t} = \ln\left(Unemp_{i,t} / Unemp_{i,t-T}\right)$	Unemployment growth rate.
Local Policy Variables	
$Tax_{i,t-T}$	The local plus regional income tax rate. This variable is divided by 100 for computational purposes.
$Exp_{i,t-T}$	Total local public expenditure per capita given by the sum of local public expenditures on culture, education, child care, family care, elderly care, environment, rescue services, subsidies to political organizations and business activities.
$Excul_{i,t-T}$	The share of local public expenditure on culture.
$Exedu_{i,t-T}$	The share of local public expenditure on education.
$Exchil_{i,t-T}$	The share of local public expenditure on child care.
$Exfam_{i,t-T}$	The share of local public expenditure on family care.
$Exeld_{i,t-T}$	The share of local public expenditure on elderly care.
Human Capital	
$Edulow_{i,t-T}$	The percentage of inhabitants between 25 and 65 years with secondary education but less than 3 years of postsecondary education.
$Eduhigh_{i,t-T}$	The percentage of inhabitants between 25 and 65 years with more than 3 years of postsecondary education.
Political representation	
$Cons_{i,t-T}$	The share of seats in the local parliament held by the Conservative party, Centre Party, Liberals and Christian Democrats.
$Herf_{i,t-T}$	A Herfindahl index measuring political stability in a municipality's parliament.
Socio-economic demographic structure	
$Age06_{i,t-T}$	The percentage of inhabitants aged 0-6.
$Age715_{i,t-T}$	The percentage of inhabitants aged 7-15.
$Age65_{i,t-T}$	The percentage of inhabitants aged ≥ 65.
$Dens_{i,t-T}$	Population density, residents per square kilometer This variable is divided by 1000 for computational purposes.
Dummy	
D	A dummy for municipalities situated in the major city areas: Stockholm, Malmö and Göteborg.

Note: Income and expenditure are normalized using the Swedish consumer price index (with 2005 as the base year). All variables are expressed in log form.

Table A2: Descriptive statistics for the overall sample (1992-2007)

Variable	Obs	Mean	Std. Dev.	Min	Max
$y_{i,t}$	1710	1.283	0.047	1.131	1.477
$m_{i,t}$	1710	0.993	0.051	0.869	1.311
$u_{i,t}$	1710	0.483	0.119	0.174	1.206
$Y_{i,t-T}$	1710	57.795	6.600	45.894	113.535
$Unemp_{i,t-T}$	1710	0.087	0.024	0.024	0.166
$Tax_{i,t-T}$	1710	30.792	1.330	25.700	34.410
$Exp_{i,t-T}$	1710	9.156	2.302	5.154	15.985
$Excul_{i,t-T}$	1710	0.078	0.021	0.068	0.298
$Exedu_{i,t-T}$	1710	0.377	0.072	0.127	0.612
$Exchil_{i,t-T}$	1710	0.180	0.049	0.127	0.390
$Exfam_{i,t-T}$	1710	0.076	0.034	0.061	0.216
$Exeld_{i,t-T}$	1710	0.339	0.098	0.081	0.706
$Edulow_{i,t-T}$	1710	0.126	0.036	0.066	0.347
$Eduhigh_{i,t-T}$	1710	0.091	0.044	0.041	0.422
$Cons_{i,t-T}$	1710	0.438	0.121	0.086	0.844
$Herf_{i,t-T}$	1710	0.285	0.058	0.160	0.492
$Age06_{i,t-T}$	1710	0.093	0.010	0.066	0.131
$Age715_{i,t-T}$	1710	0.113	0.011	0.066	0.154
$Age65_{i,t-T}$	1710	0.186	0.039	0.059	0.282
$Dens_{i,t-T}$	1710	120.985	393.711	0.276	3883.491

Table A2: Descriptive statistics for 1992

Variable	Obs	Mean	Std. Dev.	Min	Max
$y_{i,t}$	283	1.008	0.010	0.982	1.041
$m_{i,t}$	283	1.000	0.008	0.978	1.042
$u_{i,t}$	283	2.341	0.463	1.421	3.926
$Y_{i,t-T}$	283	56.715	6.001	46.604	91.273
$Unemp_{i,t-T}$	283	0.070	0.017	0.025	0.120
$Tax_{i,t-T}$	283	30.336	1.124	25.700	32.300
$Exp_{i,t-T}$	283	9.358	2.554	5.154	14.023
$Excul_{i,t-T}$	283	0.086	0.025	0.027	0.200
$Exedu_{i,t-T}$	283	0.355	0.057	0.100	0.502
$Exchil_{i,t-T}$	283	0.175	0.042	0.083	0.314
$Exfam_{i,t-T}$	283	0.066	0.029	0.013	0.189
$Exeld_{i,t-T}$	283	0.292	0.076	0.083	0.483
$Edulow_{i,t-T}$	283	0.113	0.028	0.068	0.201
$Eduhigh_{i,t-T}$	283	0.086	0.043	0.041	0.389
$Cons_{i,t-T}$	283	0.497	0.107	0.229	0.800
$Herf_{i,t-T}$	283	0.256	0.049	0.160	0.428
$Age06_{i,t-T}$	283	0.096	0.009	0.078	0.125
$Age715_{i,t-T}$	283	0.108	0.011	0.066	0.154
$Age65_{i,t-T}$	283	0.185	0.040	0.059	0.277
$Dens_{i,t-T}$	283	119.126	384.578	0.288	3655.361

Table A2: Descriptive statistics for 2007

Variable	Obs	Mean	Std. Dev.	Min	Max
$y_{i,t}$	289	1.016	0.008	0.989	1.045
$m_{i,t}$	289	1.003	0.007	0.986	1.030
$u_{i,t}$	289	0.798	0.089	0.570	1.128
$Y_{i,t-T}$	289	128.173	41.701	114.510	143.642
$Unemp_{i,t-T}$	289	0.033	0.011	0.010	0.069
$Tax_{i,t-T}$	289	31.979	0.987	28.890	34.240
$Exp_{i,t-T}$	289	13.175	1.377	9.594	17.906
$Excul_{i,t-T}$	289	0.057	0.014	0.014	0.121
$Exedu_{i,t-T}$	289	0.398	0.037	0.300	0.518
$Exchil_{i,t-T}$	289	0.156	0.039	0.074	0.303
$Exfam_{i,t-T}$	289	0.067	0.022	0.018	0.154
$Exeld_{i,t-T}$	289	0.428	0.062	0.226	0.576
$Edulow_{i,t-T}$	289	0.150	0.032	0.097	0.276
$Eduhigh_{i,t-T}$	289	0.167	0.068	0.083	0.543
$Cons_{i,t-T}$	289	0.467	0.123	0.129	0.889
$Herf_{i,t-T}$	289	0.256	0.049	0.180	0.534
$Age06_{i,t-T}$	289	0.074	0.013	0.051	0.126
$Age715_{i,t-T}$	289	0.107	0.011	0.063	0.141
$Age65_{i,t-T}$	289	0.198	0.036	0.103	0.302
$Dens_{i,t-T}$	289	131.109	442.010	0.241	4228.241

Table A3: Correlation matrix for 1992-2007

	$y_{i,t}$	$m_{i,t}$	$u_{i,t}$	Y	*Unemp*	*Tax*	*Exp*	*Excul*	*Exedu*	*Exchil*	*Exfam*	*Exeld*	*Edulow*	*Edulhigh*	*Cons*	*Herf*	*Age06*	*Age715*	*Age65*	*Dens*
$y_{i,t}$	1																			
$m_{i,t}$	0.285	1																		
$u_{i,t}$	0.108	-0.171	1																	
Y	0.165	0.410	-0.245	1																
Unemp	0.186	-0.174	0.917	-0.222	1															
Tax	-0.190	-0.346	0.297	-0.280	0.299	1														
Exp	-0.234	-0.374	0.337	-0.261	0.410	0.304	1													
Excul	-0.330	0.077	-0.167	0.206	-0.191	-0.123	-0.130	1												
Exedu	0.116	0.051	-0.290	0.212	-0.325	0.013	-0.639	-0.040	1											
Exchil	0.124	0.462	-0.360	0.690	-0.404	-0.292	-0.606	0.208	0.439	1										
Exfam	0.110	0.372	0.063	0.355	0.048	-0.107	-0.265	0.122	-0.041	0.368	1									
Exeld	0.257	-0.218	0.525	-0.339	0.527	0.461	0.091	-0.221	-0.036	-0.333	-0.202	1								
Edulow	0.279	0.554	-0.016	0.710	-0.006	-0.215	-0.241	0.215	0.081	0.600	0.431	-0.129	1							
Edulhigh	0.279	0.607	-0.169	0.705	-0.160	-0.374	-0.230	0.182	0.030	0.549	0.362	-0.251	0.887	1						
Cons	0.225	0.361	-0.353	-0.003	-0.404	-0.430	-0.284	-0.154	0.096	0.106	-0.098	-0.265	0.079	0.233	1					
Herf	-0.127	-0.218	0.264	0.081	0.311	0.234	0.197	0.155	-0.057	-0.016	0.028	0.227	-0.026	-0.139	-0.630	1				
Age06	0.163	0.256	-0.211	0.293	-0.229	-0.409	-0.210	-0.168	0.181	0.362	0.170	-0.544	0.126	0.155	0.337	-0.313	1			
Age715	0.275	-0.080	-0.098	0.107	-0.090	0.034	-0.190	-0.347	0.561	0.175	-0.106	-0.011	-0.009	-0.054	0.182	-0.139	0.398	1		
Age65	-0.074	-0.366	0.237	-0.697	0.232	0.348	0.322	-0.034	-0.333	-0.650	-0.398	0.620	-0.498	-0.445	-0.146	0.124	-0.704	-0.375	1	
Dens	0.214	0.583	-0.215	0.629	-0.229	-0.472	-0.434	0.174	0.021	0.562	0.613	-0.390	0.567	0.608	0.227	-0.149	0.304	-0.094	-0.551	1

Note: The variables are at time $t - T$

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IV

Growth and inequality: a study of Swedish municipalities

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Abstract

This paper explores the relationship between the growth rate of the average income and income inequality using data at the municipal level in Sweden for the period 1992-2007. We estimate a fixed effects panel data growth model where the within-municipality income inequality is one of the explanatory variables. Different inequality measures (Gini coefficient, top income shares, and measures of inequality in the lower and upper ends of the income distribution) are also examined. We find a positive and significant relationship between income growth and income inequality, measured as the Gini coefficient and top income shares, respectively. In addition, while inequality at the upper end of the income distribution is positively associated with the income growth rate, inequality at the lower end of the income distribution seems to be negatively related to the growth rate. Our findings also suggest that increased income inequality enhances growth more in municipalities with a high level of average income than in those with a low level of average income.

JEL classification: R11, D31, D63, O47

Keywords: Growth, inequality, Gini coefficient, panel data model, municipalities

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1. Introduction

The effect of income inequality on economic growth has attracted much interest among economists, as well as in popular debate (Stiglitz, 2012). A key issue in the literature refers to the possible trade-off between equality and economic growth. With the expansion of the welfare state, Sweden experienced a significant decrease in income inequality, measured in terms of taxable income, after World War II (Lindbeck, 1997) followed by a slight increase in the last decades. It is not surprising, therefore, that Sweden, thanks to its tradition as an egalitarian society, has attracted interest from scholars (Roine and Waldenström, 2008). Recent research based on county level data for Sweden suggests that increased income inequality leads to increased income growth, supporting the idea of a trade-off between growth and equality (Nahum, 2005; Gromark and Petersson, 2010).

Economists typically address the relationship between growth and inequality by adding a measure of income inequality as an additional explanatory variable in a growth regression model. Based on this approach, and by focusing on cross-country data, an extensive literature has explored how the distribution of income affects the growth rate of an economy's gross domestic product (GDP)[1]. Several earlier studies on the growth-inequality relationship examine a single cross-section of countries and typically find a negative and significant relationship between GDP growth and income inequality (Persson and Tabellini, 1994; Alesina and Perotti, 1996). Other studies, such as Forbes (2000), instead report a significant and positive relationship based on cross-country panel data using fixed effects estimation. Similarly, Partridge (1997) finds a positive correlation between GDP growth and the Gini coefficient based on panel data for the US states. Li and Zou (1998) find that the relationship between income inequality and economic growth becomes significantly positive when using panel data (along with appropriate methods for handling such data), while a negative relationship is typically found for single cross-section data. There is also evidence suggesting that the relationship between growth and inequality may vary systematically between countries: Barro (2008) finds that inequality appears to encourage growth only within rich countries and to slow down growth in poorer countries. His paper covers data from 1960 to 1990 for 84 countries.

In the literature, at least three arguments have been put forward for a positive relationship between growth and inequality. The first argument is that inequality enhances growth through greater investment opportunities. This argument assumes that investments are characterised by large set-up costs and that there exists credit-market imperfections, i.e. there are limits to borrowing money. In

[1] The opposite causality is the Kuznets' Hypothesis, which postulates that income inequality first rises and then falls during the course of economic development (Kuznets, 1955).

that set-up, an initial amount of personal wealth is needed to be able to invest in a project. With an unequal wealth distribution, therefore, one may find more wealthy people who are able to invest and initiate new industrial activity, which would enhance economic growth (Aghion *et al.*, 1999). The second argument is based on the idea that individual savings rates increase with the level of income. Therefore, a redistribution of resources from rich to poor people may lower the aggregate rate of savings. Through this channel, increased inequality tends to increase investment (Aghion *et al.*, 1999). This effect is primarily relevant for closed economies, where domestic investments equal the domestic savings. Finally, the last argument is that equal wages would discourage people from exerting maximal effort. Instead, in a society with an unequal wage structure, workers would have an incentive to exert effort to receive a higher wage (Aghion *et al.*, 1999; Voitchovsky, 2005). In a similar way, there are also arguments for a negative relationship between growth and inequality. Accordingly, a more equal income distribution may enhance growth by reducing corruption, crime, and social unrest, and by alleviating credit constraints, allow for investment in human and physical capital (Alesina and Rodrik, 1994; Persson and Tabellini, 1994; Alesina and Perotti, 1996; Aghion *et al.*, 1999; Barro, 2000).

The present study attempts to explore the relationship between average income growth and income inequality at the municipal level in Sweden from 1992 to 2007 using different measures of income inequality. This is accomplished by estimating a growth-inequality model to test whether and how income inequality affects income growth, conditional on a set of control variables. Data on income inequality refer to population shares in different income classes for each municipality and year, which can then be used to calculate measures of income inequality such as the Gini coefficient and different top income shares. Following Voitchovsky (2005), we also estimate a growth model including measures of inequality both at the lower and upper ends of the income distribution. Furthermore, we examine if the relationship between income growth and income inequality varies with the income level by allowing for an interaction effect between the Gini coefficient and the level of average gross income (see Barro, 2000, 2008, for a similar approach based on cross-country panel data).

To our knowledge, there are no earlier studies dealing with the relationship between income inequality and growth at the municipal level in Sweden. Instead, previous studies are based on data at the county level (Nahum, 2005) or labour market regions (Rooth and Stenberg, 2012) whereas our study focuses on a panel data set covering 283 municipalities. Nahum (2005) explores the relationship between growth and inequality measured by the Gini coefficient using panel data at the county level from 1960 to 2000. She estimates the effects of inequality in growth regressions with 1, 3, 5 and 10-year growth periods, where the growth rate depends on explanatory variables measured at the beginning of each growth period. Using fixed effects panel 2SLS estimations, she finds a positive effect of inequality on 1 to 5-year economic growth rates (when significant), whereas the corresponding effect based on 10-year growth periods is typically not significant.

Rooth and Stenberg (2012) analyse the relationship between growth and income inequality based on data for 72 labour market regions in Sweden during the period 1990-2006. The labour market regions are based on commuting patterns of the labour force, which have been constructed by Statistics Sweden, collapsing 289 municipalities into 72 regional labour markets. The model is estimated by using cross-sectional OLS, panel (pool) OLS, panel with region specific fixed effects and system GMM[2]. The main results show that increased overall inequality (through an increase in the Gini coefficient) and increased inequality in the upper end of the income distribution (measured by the ratio between the 90th and the 50th income percentiles) lead to increased growth, while they find no evidence that inequality in the lower end of the income distribution (measured by the ratio between the 50th and the 10th income percentiles) affects the income growth rate.

However, the results presented in Rooth and Stenberg (2012) are sensitive to the choice of estimation method and sample size. In a cross-sectional estimation, which includes one observation per region, the Gini coefficient is positively related to growth, while in a pooled OLS model the Gini coefficient has a negative effect on growth when combined with the measure of inequality in the upper end of the distribution. Also, the effect of the Gini coefficient changes from being positive to negative when the time span of the model changes from 1990-2006 to 1994-2006. Finally, when a system GMM model is estimated for the period 1990-2006, the Gini coefficient has a negative effect on growth, while for the period 1994-2006 the Gini coefficient has a positive effect if combined with the measure of inequality in the upper end of the income distribution.

Compared to other studies based on Swedish data (Nahum, 2005; Rooth and Stenberg, 2012), we use a broader set of inequality measures to study the relationship between income growth and income inequality. We consider the Gini coefficient (as a measure of the overall inequality) as well as the income shares of the 25%, 15%, 10%, 5% and 1% top income earners (not included in the above studies). Moreover, we follow Rooth and Stenberg (2012) by analysing the effects of inequality at the lower end of the income distribution, measured by the ratio between the 50th and 10th income percentiles (referred to as 50/10), and inequality at the upper end of the income

[2] Cross-sectional regressions are based on average annual growth 1990-2006 and 1994-2006, Pool and system GMM used data measured in 4-year intervals.

distribution measured by the ratio between the 90th and the 75th income percentiles (referred to as 90/75). This allows us to cover most of the inequality measures used in the literature. A further important contribution is to examine the simultaneous effects of different measures of income inequality, i.e. the joint effects of the Gini coefficient and the 90/75 and 50/10 measures, in terms of municipal income growth, which facilitates comparison with the results for labour market regions presented by Rooth and Stenberg (2012).

Our data covers 15 years, from 1992 to 2007 and are collected from the same source, Statistics Sweden (SCB), in a uniform manner across municipalities to reasonably ensure comparability and minimise measurement errors. This is the most recent and longest period from which we could find comparable data at the municipal level, that is, all variables are measured in a consistent way over time. An advantage with our data is the high degree of homogeneity in terms of democratic functions, public transfer systems, educational systems, and labour market institutions across municipalities in the same country; a similar argument was made by Rooth and Stenberg (2012) in the context of labor market regions. In fact, a problem with cross-country studies is that the data on income statistics may differ significantly between countries regarding both the quality and definitions (Pardo-Beltrán, 2002). Moreover, countries may also differ in the "level of democracy, human rights, type of economy, education system etc., which does not make it reasonable to expect that one single model holds for all countries" (Nahum, 2005). These issues could be less severe when using data for a single country, where much of the institutions are the same.

The paper is structured as follows: Section 2 provides the empirical framework where the model is presented together with data. In Section 3, the results are presented and discussed. The last section, Section 4, provides the conclusions.

2. Empirical framework

This section presents a regression model for studying the relationship between income growth and income inequality, conditional on a set of other control variables that may affect the growth rate. The analysis follows the 5-year panel data growth model examined in several recent papers such as Li and Zou (1998), Voitchovsky (2005) and Barro (2008), as well as the set-up and explanatory variables used in earlier studies on economic growth at the regional/local level in Sweden (e.g. Aronsson *et al.*, 2001; Lundberg, 2003; and Nahum, 2005).

The basic regression model, with income inequality measured by the Gini coefficient, is given by the following equation, where the dependent variable, $y_{i,t}$, denotes the growth rate of the average income among the municipal residents between periods $t-T$ and t:

$$
\begin{aligned}
y_{i,t} = {} & \alpha_i + \beta Y_{i,t-T} + \delta_1 G_{i,t-T} + \delta_2 Tax_{i,t-T} + \delta_3 Exedu_{i,t-T} + \delta_4 Exfam_{i,t-T} \\
& + \delta_5 Exchil_{i,t-T} + \delta_6 Exeld_{i,t-T} + \delta_7 Exp_{i,t-T} + \delta_8 Edu_{i,t-T} + \delta_9 Herf_{i,t-T} \\
& + \delta_{10} Cons_{i,t-T} + \delta_{11} Age65_{i,t-T} + \delta_{12} (G \cdot Y)_{i,t-T} + \varepsilon_{i,t}
\end{aligned} \tag{1}
$$

where $i = 1,...,N$ refers to cross-section unit (municipality), while α_i, β, and the δ's are parameters to be estimated. The variable $\varepsilon_{i,t}$ represents an independently and identically distributed error term. In the estimations, we use $T = 5$ years[3]. Taking five-year averages will reduce the short-run fluctuations and therefore the influence of the economic cycle. The five-year time lag also facilitates comparison of our analysis with other similar studies (e.g. Barro, 2000; Nahum, 2005[4]; Voitchovsky, 2005; Rooth and Stenberg, 2012), which have also used the same time lag. The variables used in the estimation include a broad spectrum of potential determinants of average income growth. The list and detailed definitions of the variables are presented in Table 1.

The initial level of average income, $Y_{i,t-T}$, i.e. the average income at the beginning of each growth-period, allows us to control for the relationship between the income level and the subsequent growth, and also address the question of conditional convergence. The latter means that each municipality is converging to its own steady state (see, for example, Barro and Sala-i-Martin, 1992, and Mankiw *et al.,* 1992). Previous studies based on Swedish data have found that regions tend to converge over time (either unconditionally, as in Persson, 1997, or conditionally, as in Aronsson *et al.*, 2001; Nahum, 2005). As a consequence, we control for this mechanism through the initial level of average income in the municipality.

The variable $G_{i,t-T}$ denotes the Gini coefficient for municipality *i* in period *t-T*. As indicated above, we will also consider other measures of income inequality such as different top income shares as well as the measures of inequality in the upper and lower ends of the income distribution. This will

[3] As mentioned in section 3 below, we have also considered 3- and 4-year growth periods.
[4] As mentioned above, she also estimated models based on longer growth periods.

Table 1: Description of the variables

Variable	Description
$y_{i,t}$	Average income growth $y_{i,t} = \ln(Y_{i,t} / Y_{i,t-T})$ where $Y_{i,t}$ is the average income level (thousands of SEK) measured for the population aged 20 and above in municipality *i* between period *t-T* and *t*.
$Y_{i,t-T}$	Log of average income (thousands of SEK) among those aged 20+.
$G_{i,t-T}$	The Gini coefficient, which has been calculated using data on income distribution for the population aged 20 and above at the municipal level. For each year and municipality, these data contain information about the population in different intervals of taxable personal income[5]. The formula for the Gini coefficient is presented in Appendix A1. G is equal to 1 when the inequality is at its maximum and zero with an equal distribution.
$Top\%_{i,t-T}$	Income shares of top 25%, 15%, 10%, 5% and 1% income earners (our elaboration) i.e. the sum of income for people in the richest group divided by the sum of income over all residents.
$90/75_{i,t-T}$	The income of the 90th percentile divided by the income of the 75th percentile (our elaboration).
$50/10_{i,t-T}$	The income of the 50th percentile divided by the income of the 10th percentile (our elaboration).
$Tax_{i,t-T}$	Local income tax rate given by the sum of the municipal and county income tax rates.
$Exedu_{i,t-T}$	The share of expenditure on education.
$Exchil_{i,t-T}$	The share of expenditure on child care.
$Exfam_{i,t-T}$	The share of expenditure on family care.
$Exeld_{i,t-T}$	The share of expenditure on elderly care.
$Exp_{i,t-T}$	Per capita total expenditure (SEK) given by the sum of expenditure on culture, education, childcare, family care, and on elderly care divided by the population (for computational reasons, divided by 1000).
$Edu_{i,t-T}$	The share of inhabitants with at least three years of university education.
$Herf_{i,t-T}$	Herfindahl index measuring the political stability in the municipality parliament. The Herfindahl index[6] is measured by the sum of the squared shares of seats in the municipal council occupied by the Conservative Party, Centre Party, Liberal Party, Christian Democrats, Green Party, Social Democrats, Left Party and other parties.
$Cons_{i,t-T}$	The share of seats in the local parliament held by the Conservative Party, Centre Party, Liberals and Christian Democrats.
$Age65_{i,t-T}$	The share of inhabitants aged 65 and above.
$Dens_{i,t-T}$	Population density, measured as the number of residents per square kilometre.

Note: Income and Expenditure are adjusted by Swedish CPI (2005 is the base year). Measures of income inequality are also based on real income.

be described in greater detail below.

As suggested in earlier studies, the growth rate of average income also depends on local and national policy decisions (see Glaeser *et al.*, 1995; Helms, 1995; Aronsson *et al.,* 2001). The effects of local

[5] The best available data are based on taxable personal income. We also use taxable personal income, like Nahum (2005), for reasons of comparison. Taxable income was also used by Persson and Tabellini (1994) and Perotti (1996).

[6] The Herfindahl index is defined as $\sum_{p=1}^{P} SH_p^2$ where SH_p is the share of seats from party *p*. The Herfindahl index ranges from a minimum of $1/P$ to a maximum of 1 if a single party holds all seats in the local council, where P is the number of different parties.

policy variables are captured by the parameters $\delta_2,...,\delta_7$. The local income tax rate, Tax, given by the sum of the municipal and county income tax rates, serves to capture differences in the tax-wedge between municipalities. Helms (1985) found a negative effect of the income tax rate on economic growth based on data from 1965 for 48 states, whereas the effect of the local income tax rate is typically insignificant in earlier studies of income growth at the regional and municipal level in Sweden (Aronsson *et al.,* 2001; Lundberg, 2003).

We control for the effects of the public expenditure pattern via different municipal public expenditure shares: $Exedu$, the share of local public expenditure on education; $Exfam$, the share of local public expenditure on family care; $Exchild$, the share of local public expenditure on childcare; and $Exeld$, the share of local public expenditure on elderly care. The variable Exp denotes the per capita total expenditure (in SEK), reflecting public spending on education, family care, child care, elderly care and culture. The effect of local public expenditure per capita may capture effects of productive expenditure as well as disincentives associated with revenue collection. Aronsson *et al.* (2001) find that local public expenditure per capita does not significantly affect the economic growth based on income data at the county level, whereas Lundberg (2003) finds that the local public expenditure per capita has a negative impact on average income growth within the major city areas (Stockholm, Malmö and Goteborg) based on income data at municipal level.

Human capital is typically expected to enhance economic growth. In the present study, we try to control for human capital by including the variable Edu, i.e. the share of the population between 25 and 65 with three years or more of university education, among the regressors. Several earlier studies have used the same definition of human capital (Partridge, 1997; Aronsson *et al.*, 2001; Panizza, 2002; Lundberg, 2003; Nahum, 2005).

Average income growth may also depend on factors related to the political preferences and stability of the local council. As such, and following Lundberg (2003), we include the Herfindahl index, $Herf$, and a dummy variable for liberal/conservative party majority, $Cons$, among the regressors. The Herfindahl index is an indicator of political stability in the sense that an increase in this index means less fragmentation in the municipal parliament. As such, a decrease in the Herfindahl-index is expected to reduce the income growth rate, due to that a fragmented parliament makes it difficult to reach long-term agreements (see Roubini and Sachs, 1989a, 1989b; Alesina and Perotti, 1995; and Alesina *et al.*, 1996). On the other hand, we have no prior expectations regarding the effect of the share of seats in the local parliament held by the liberal and conservative parties. Finally, we also control for the demographic composition as represented by the share of the population aged 65 and above, ($Age65$), and population density, $Dens$; see, for example, Perrotti (1996), Forbes (2000), Panizza (2002), and Nahum (2005).

2.1 Descriptive Statistics

In this section, we discuss the descriptive statistics for the main variables as displayed in Table A1, Table A2, Figure A2 and Figure A3 in the appendix. The dataset originates from Statistics Sweden and covers 283 out of 290 municipalities[7], during the period 1992-2007. The measures of income inequality are calculated by using information about the distribution of personal taxable income in each municipality and year. More specifically, the population in each municipality is divided into 18 income classes; by assuming that each individual has the average income in his/her income class, it is then possible to calculate the Gini coefficient and the top income shares (top 25%, 20%, 15%, 10%, 5%, and 1%). The inequality measures also include the top and bottom end inequality. More precisely, top end inequality is measured by the income percentile ratio 90/75, and bottom end inequality is measured by the income percentile ratio 50/10.

Looking at Table A1, income growth, y, is on average 2% for the entire period, and the standard deviation of income growth rate is larger within, than between, municipalities. The variability of the average income, Y, measured by the standard deviation within municipalities is almost the same as the standard deviation between municipalities. Municipalities in the county of Stockholm such as Danderyd, Lidingö, Täby, Sollentuna, and Vaxholm, appear to have the highest average income every year. The average Gini coefficient is 0.16 for the overall period with a standard deviation of 0.05; the standard deviation is larger between (0.045), than within (0.031), municipalities. By comparing extremes, we have observed that the largest average income is in the municipality of Danderyd in 2007 (439.6 thousand SEK), which also has the largest Gini coefficient (0.55) in the same year. The smallest average income (115 thousand SEK) can be found in Bjurholm in 1992, while the lowest Gini (0.07) is found in Överkalix in 1992. Thus, in line with earlier studies, most of the variation in the Gini coefficients at the municipal level is cross-sectional rather than temporal (e.g. Deininger and Squire, 1996; Partridge, 2005; Voitchovsky, 2005).

Turning to the other measures of inequality to be used below, notice first that the municipalities with high top income shares typically are those situated in the Stockholm region. The municipality with the highest top income shares is Danderyd. The ratio 90/75 seems to show some variation

[7] The municipalities Nykvarn, Knivsta, Gnesta, Trosa, Gotland, Bollebygd, and Lekeberg were excluded from our analysis because borders were changed during 1992-2007.

both between and within municipalities during 1992-2007. On the other hand, the ratio 50/10 varies greatly, both over time and across municipalities. The inequality measures also have an increasing trend towards the end of the sample period (see Fig. A2 in the appendix). The same pattern can also be noticed for average income in some municipalities (see Fig. A3 in the appendix). The similarity of the trends suggests that there is some type of relation between inequality and economic growth at the municipal level.

2.2 The endogeneity issue

The estimation of the effect of income inequality on economic growth raises some problems of endogeneity. First, there is a correlation between $Y_{i,t-T}$ on the right side of (1) and the fixed effect. Another potential endogeneity problem is associated with the income inequality measures; the estimated effect of income inequality on the income growth rate might be biased by a correlation between the measure of inequality and the error term. If the RHS variables are endogenous and thus correlated with the error term, the OLS/FE coefficient estimate is biased and inconsistent[8]. Yet, the fact that the RHS variables are dated at the beginning of the growth period naturally minimises the problem of endogeneity.

A number of authors cope with these problems by using 3SLS (Barro, 2000, 2008) or System-GMM (Castelló-Climent, 2004). The GMM procedure of first-differencing, using lags as instruments, requires more data points than in our case study. Yet, based on the Monte Carlo simulations in Judson and Owen (1999), it is not necessarily incorrect to treat income variables as exogenous: they show that for panels with a short-time dimension (as in our case), the bias of the coefficient on the lagged dependent variable can be significant, while the bias of estimates of the other coefficients may still be minor[9]. Nevertheless, the endogeneity problem may persist.

The inequality measured in period $t-T$ may be correlated with the same unobserved characteristics as the income growth rate between periods $t-T$ and t. Some of the previous works (Alesina and Rodrik, 1994; Nahum, 2005)[10] discuss whether the Gini coefficient should be considered as an exogenous or an endogenous variable. Alesina and Rodrik (1994) assume that the Gini coefficient is

[8] The estimation of the growth-inequality link is plagued by omitted variable bias, where unobserved time invariant characteristics may correlate both with the lagged dependent variable and the fixed effect. Nonetheless, the possible effects of unobserved attributes in a single country context are most likely not as important as in a multi-country framework, since institutions, culture and norms are largely similar across municipalities in Sweden. Therefore, omitted variable bias should be alleviated as municipalities have similar redistributive policies and institutions (Rooth and Stenberg, 2012).

[9] We have also estimated our model considering income as exogenous. The signs and significance of the coefficients were consistent with the results presented in Table 2, so we have not presented these results.

[10] They argue that income inequality and income growth might be jointly determined.

an endogenous variable and use 2SLS to estimate the growth equation. Nahum (2005) addresses the endogeneity problem by using 2SLS, where the Gini coefficient is regressed on variables reflecting the age structure of the population in the first stage. She also compares the second stage results for the growth equation with those that follow if the Gini coefficient is treated as exogenous, and finds no important differences in the results. Assa (2012) also examines the relationship between economic growth and income inequality using data for 141 countries between 1992 and 2005. He estimates a 2SLS regression and uses primary school enrolment as an instrumental variable for the Gini coefficient.

So, following the works mentioned above, we estimate the growth regression of our panel data model with fixed effects and instrument $Y_{i,t-T}$, with a one period income lag $Y_{i,t-T-1}$, (Green, 2011). In a way similar to Nahum (2005), who also addresses the endogeneity of income inequality, we estimate our model using 2SLS to instrument inequality, where the first stage regression for the Gini coefficient, top income shares, 90/75 and 50/10, respectively, uses the share of age groups (such as Age24_44, Age45_64, Age65) in the adult population as explanatory variables. The age structure has a direct effect on the income distribution, since it captures the shares of the population in various parts of their life cycle.

3. Empirical results

As indicated above, we have performed first stage regressions with the inequality measures as dependent variables. We investigate the strength our instruments using the first stage F-statistic as suggested by Bound *et al.* (1995) and Steiger and Stock (1997). Based on the F-test, we find that the instruments have significant effects in each of the first stage regressions[11]. Our F-statistics exceed the rule of thumb value[12] of 10; this implies that our instruments offer a good explanation of the variation in the endogenous variables[13].

The results of estimating model (1) are summarised in Tables 2 and 3. In Table 2, we report in column (i) the relationship between the average income growth rate and the Gini coefficient without

[11] To save space, the results from the first stage regressions are not presented here, although available from the author upon request.
[12] Stanger and Stock (1997) provide a theoretical analysis of the properties of the IV estimator and provide some guidlines about how large the F-statistic should be for the IV estimator to be a valid instrument. For the rule of thumb, see also Stock and Watson (2005, Ch.10).
[13] We also test if the inequality measures are endogenous by applying the Durbin–Wu–Hausman test. See Hausman (1978), and also Durbin (1954) and Wu (1973). See also Green (2011). We have endogeneity if the 2SLS estimates differ significantly from the OLS estimates. We reject the null hypothesis, so we have an endogenenity issue.

any interaction effect between the Gini coefficient and the level of average income. In column (ii), we follow Barro (2000, 2008) and Forbes (2000) and add such an interaction effect to examine whether the income growth rate in municipalities with a high level of average income responds to income inequality in a different way than the income growth rate in municipalities with a lower level of average income. In columns (iii) and (iv) and in columns (v) and (vi), respectively, we carry out the same analysis with the modification that income inequality is measured by top income shares (top 25 and top 15) instead of by the Gini coefficient. Results for other top income shares can be found in the appendix, Table A4. Finally, in Table 3, we present the estimation results for a more comprehensive model, which contains an overall measure of income inequality (given by the Gini coefficient) as well as measures of inequality in the lower (50/10) and upper (90/75) ends of income distribution.

Table 1 shows that the effect of the Gini coefficient and top income shares, respectively, on the income growth rate is always positive and significant in all columns. These results are in line with Nahum (2005), who finds that an increase in the Gini coefficient leads to an increase in the average income growth rate at the country level in Sweden. Although we use a different inequality measure for the upper part of the income distribution, our results are also consistent with Voitchovsky (2005), who finds that inequality in the upper part of the income distribution has a positive effect on the growth rate[14].

Furthermore, notice that the parameter of the interaction variable $(G \cdot Y)$ is significant and has a positive sign, suggesting that the relationship between the growth rate and the Gini coefficient is stronger in municipalities with a high level of average income than in municipalities with a low level of average income, *ceteris paribus*. This result is also confirmed for the specification with $(Y \cdot Top\%)$. As such, it appears as if the trade-off between growth and equality is stronger in rich municipalities, although the trade-off is also present in poor municipalities. The positive effect of the interaction variable is in line with Barro's (2000, 2008) earlier findings on cross-country data and is also found by Forbes' (2000).

Turning to the effects of the other regressors, the results show that the initial level of real average income, Y, has a negative and significant effect on the income growth rate, suggesting conditional

[14] As mentioned in the introduction, the positive relationship between growth and income inequality is also consistent with previous studies based on cross-country panel data (Forbes (2000) and Li and Zou (1998)), as well as results presented in Partridge (1997) based on income data for the US states. On the other hand, based on similar data for the period 1940-1990, Panizza (2002) found evidence of a negative relationship between income inequality and economic growth.

convergence in the sense that municipalities with a low level of initial income tend to grow faster than municipalities with high income levels, conditional on the other explanatory variables. A similar result has been found in studies using data on US states (Barro and Sala-i-Martin, 1992; 1995) and studies based on data for Swedish counties and municipalities (Aronsson *et al.,* 2001; Lundberg, 2003).

Table 2: Fixed Effects estimations

	Gini[15]		Top25		Top15	
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
	ln(y)					
Constant	-0.040 (-0.76)	-0.005 (-0.09)	0.322*** (5.23)	0.404*** (6.53)	0.335*** (5.66)	0.404*** (6.83)
G_{it-T} or $Top\%_{i,t-T}$	0.517*** (10.80)	0.564*** (11.80)	0.703*** (8.88)	0.799*** (10.08)	0.731*** (10.63)	0.800*** (11.65)
$Y_{i,t-T}$	-0.215*** (-6.00)	-0.624*** (-9.77)	-0.284*** (-8.21)	-0.701*** (-10.97)	-0.274*** (-8.00)	-0.690*** (-10.83)
$Tax_{i,t-T}$	-0.059 (-0.90)	-0.088 (-1.35)	-0.0316 (-0.48)	-0.061 (-0.92)	-0.067 (-1.01)	-0.097 (-1.48)
$Exedu_{i,t-T}$	0.246*** (4.32)	0.231*** (4.10)	0.236*** (4.10)	0.224*** (3.94)	0.246*** (4.31)	0.231*** (4.09)
$Exfam_{i,t-T}$	0.681*** (11.03)	0.667*** (10.90)	0.651*** (10.51)	0.636*** (10.37)	0.652*** (10.60)	0.635*** (10.42)
$Exchil_{i,t-T}$	0.362*** (5.76)	0.317*** (5.08)	0.383*** (6.05)	0.341*** (5.42)	0.366*** (5.82)	0.321*** (5.14)
$Exeld_{i,t-T}$	0.399*** (7.68)	0.382*** (7.41)	0.382*** (7.29)	0.367*** (7.08)	0.388*** (7.49)	0.370*** (7.21)
$Exp_{i,t-T}$	-0.086*** (-14.16)	-0.093*** (-15.29)	-0.091*** (-14.59)	-0.097*** (-15.61)	-2.267*** (-21.31)	-0.093*** (-14.97)
$Edu_{i,tT}$	2.403*** (21.64)	2.300*** (20.77)	2.165*** (20.51)	2.052*** (19.46)	0.086*** (13.87)	2.153*** (20.25)
$Herf_{i,t-T}$	0.104*** (5.79)	0.117*** (6.58)	0.101*** (5.61)	0.114*** (6.36)	0.107*** (5.98)	0.121*** (6.78)
$Cons_{i,t-T}$	-0.328*** (-24.74)	-0.327*** (-24.91)	-0.327*** (-24.50)	-0.326*** (-24.71)	-0.325*** (-24.57)	-0.324*** (-24.73)
$Age65_{i,t-T}$	-0.048 (-0.52)	0.001 (0.01)	0.018 (0.19)	0.078 (0.82)	0.052 (0.55)	0.111 (1.18)
$Dens_{i,t-T}$	-0.089* (-2.55)	-0.074* (-2.12)	-0.148*** (-4.31)	-0.135*** (-3.99)	-0.127*** (-3.72)	-0.115*** (-3.39)
$(G \cdot Y)_{i,t-T}$ or $(Top\% \cdot Y)_{i,t-T}$		0.895*** (7.70)		0.908*** (7.73)		0.898*** (7.71)
N of observations	3113	3113	3113	3113	3113	3113

Note: t-values in parentheses. Significance level: "*"p<0.05, "**", p<0.01, "***" p<0.001

[15] We also estimated this model using three lags. We noticed that the sign and significance of the effect of the Gini coefficient did not change, and the changes in magnitudes and significance levels of the other regression coefficients were very small, so we did not report these results.

The local income tax rate, Tax, is negatively associated with the income growth rate, although the estimated coefficient is insignificant. In addition, the initial level of local public expenditure per capita has a negative and significant effect on the subsequent income growth rate. These results are comparable to some of the findings presented in earlier studies based on Swedish data (Aronsson *et al.*, 2001; Lundberg, 2003) discussed above. Yet, notice also that the effects of the local policy variables are not robust across different versions of the model: in Table 3 the estimated effect of the local income tax rate is negative and significant, while the local public expenditure per capita has a positive and significant effect on the income growth rate. It is difficult to give interpretations of these results as local councils were not required to balance their budget each year during this period[16]. Consequently, local government expenditure and income tax rates may not solely reflect the current service level and cost for taxpayers. The shares of local public expenditure on education, family, children, and elderly are all significantly and positively correlated with income growth in all models.

It is expected that economic growth and human capital are positively correlated. In accordance with these expectations, we find that the share of the population with a university education of three or more years, Edu, is positively correlated with the growth rate in all columns. This result is consistent with theory, and indicates the importance of human capital for municipal income growth. The Herfindahl index, $Herf$, has a positive and significant effect in the growth equation. Therefore, a more fragmented municipal parliament (as represented by a lower index value) tends to lower the income growth rate, *ceteris paribus*. Although expected, this result is in contrast to what is found in Lundberg (2003), where the coefficient of $Herf$ is negative and significant. The results also indicate that municipalities with a liberal and/or conservative majority, $Cons$, in the local parliament are characterised by a lower growth rate than other municipalities, *ceteris paribus*. Finally, the effect of the share on inhabitants aged 65 and above, $Age65$, in the growth equation is insignificant, whereas the population density, $Dens$, has a significant and negative effect on the subsequent growth rate.

Returning to the effects of income inequality, we perform a robustness check, as in Barro (2000), by dividing the sample into one group with high-income municipalities and another with low-income municipalities, where the sample mean constitutes the break point. In this case we can estimate two separate coefficients. The results of our investigation, based on fixed effects models, are presented in the appendix in Table A5. As we can see, the coefficients of G are positive and significant for

[16] Some other studies find a negative relationship between public expenditure and economic growth (Barro, 1991; Gwartney, Lawson and Holcombe, 1998). These studies look at spending beyond certain core functions and conclude that growth will be retarded if public expenditure is too high.

both rich and poor municipalities although they differ in term of magnitude; the coefficient of G for rich municipalities is larger than the coefficient for the poor municipalities, which is consistent with the results presented above. We also test for significant differences between the rich and poor municipalities with respect to the effect of the Gini coefficient, by including all observations in the same regression and using a dummy variable to indicate the rich municipalities. We reject the hypothesis that the effect of the Gini coefficient is the same in both types of municipalities (*t*-value 6.98; 95% CI: 1.132 to 2.016), and conclude that the effect is larger for rich municipalities than for poor municipalities.

The next step is to follow Voitchovsky (2005) and test whether top end and bottom end inequality affect income growth in different ways, and also examine the simultaneous effects of the Gini coefficient and the 90/75 and 50/10 ratios. These results are presented in Table 3, where the 90/75 ratio is found to have a positive and significant effect on the income growth rate, while the effect of the 50/10 ratio is negative and significant. A joint test shows that the 90/75 and 50/10 measures jointly contribute to the municipal income growth rate. When we consider the three inequality measures (Gini, 90/75 and 50/10) simultaneously, our results show that both the Gini and the 90/75 ratio have positive and significant effects on the income growth rate, while the 50/10 ratio has a negative effect. The effects of the Gini coefficient and the 90/75 ratio may reflect some of the mechanisms discussed in Section 1. More inequality at the top may indicate that more individuals are wealthy enough to carry out investment projects with high fixed costs. It may also imply stronger incentives to undertake effort. Moreover a further explanation of this result is provided by Galor and Tsiddon, (1997) and Hassler and Mora, (2000) who argue that a concentration of talented and skilled individuals with a high income share (in advanced technology sectors) may be conducive to technological progress, and therefore to growth.

Inequality outside the top range of the income distribution, after controlling for the effect of the Gini coefficient, seems to have a negative effect on the income growth rate and may be associated with channels such as credit constraints and investment in human capital, as well as increased crime and insecurity. As a result of limited funds, some individuals might not be able to exploit their skills and talent, and fewer productive investments will be undertaken (Galor and Moav, 2004). Still, Persson and Tabellini (1994) argue that the negative effect of inequality on economic growth coexists with imperfect credits markets; poor people may be unable to invest in their human and physical capital with negative consequences for growth.

Table 3: Fixed Effects estimations with top and bottom end inequality

	Top only	Bottom only	G and Top	G and bottom	Top and Bottom	G and Top and bottom
	ln(y)					
Constant	-0.206*** (-4.08)	-0.139** (-3.14)	-0.206*** (-4.13)	-0.141** (-3.21)	-0.198*** (-3.94)	-0.201*** (-4.02)
$G_{i,t-T}$			0.331*** (6.31)	0.314*** (5.93)		0.314*** (5.95)
$90/75_{i,t-T}$	0.050* (2.51)		0.0497* (2.55)		0.0484* (2.47)	0.0490* (2.51)
$50/10_{i,t-T}$		-0.006** (-3.12)		-0.005* (-2.30)	-0.006** (-3.08)	-0.005* (-2.26)
$Y_{i,t-T}$	-0.362*** (-34.31)	-0.367*** (-34.27)	-0.437*** (-27.44)	-0.438*** (-27.44)	-0.368*** (-34.33)	-0.438*** (-27.50)
$Tax_{i,t-T}$	-0.333*** (-7.04)	-0.340*** (-7.19)	-0.378*** (-7.95)	-0.381*** (-8.03)	-0.335*** (-7.09)	-0.377*** (-7.94)
$Exedu_{i,t-T}$	0.255*** (5.78)	0.256*** (5.81)	0.237*** (5.41)	0.240*** (5.47)	0.253*** (5.74)	0.237*** (5.40)
$Exfam_{i,t-T}$	0.275*** (5.72)	0.272*** (5.65)	0.250*** (5.22)	0.250*** (5.22)	0.268*** (5.57)	0.246*** (5.14)
$Exchil_{i,t-T}$	0.439*** (9.11)	0.436*** (9.05)	0.432*** (9.04)	0.432*** (9.03)	0.431*** (8.96)	0.427*** (8.93)
$Exeld_{i,t-T}$	0.256*** (6.33)	0.257*** (6.36)	0.224*** (5.53)	0.227*** (5.62)	0.253*** (6.26)	0.223*** (5.52)
$Exp_{i,t-T}$	0.024*** (4.25)	0.024*** (4.40)	0.027*** (4.83)	0.027*** (4.91)	0.024*** (4.38)	0.027*** (4.90)
$Edu_{i,t-T}$	1.549*** (19.44)	1.535*** (19.23)	1.363*** (16.14)	1.362*** (16.13)	1.534*** (19.24)	1.361*** (16.13)
$Herf_{i,t-T}$	0.026* (2.04)	0.025 (1.93)	0.020 (1.56)	0.018 (1.46)	0.026* (2.07)	0.020 (1.60)
$Cons_{i,t-T}$	-0.156*** (-15.75)	-0.157*** (-15.89)	-0.146*** (-14.72)	-0.148*** (-14.90)	-0.156*** (-15.75)	-0.147*** (-14.76)
$Age65_{i,t-T}$	-0.467*** (-5.94)	-0.471*** (-6.00)	-0.615*** (-7.55)	-0.613*** (-7.52)	-0.461*** (-5.88)	-0.603*** (-7.40)
$Dens_{i,t-T}$	-0.081 (-1.11)	-0.080 (-1.08)	-0.110 (-1.51)	-0.107 (-1.47)	-0.08 (-1.07)	-0.107 (-1.46)
N of observations	3113	3113	3113	3113	3113	3113

Note: t-values in parentheses. Significance level: "*"p<0.05, "**", p<0.01, "***" p<0.001

In summary, we can see that all specifications suggest that the ratio 50/10 has a negative effect on growth and the ratio 90/75 has a positive effect on growth in Sweden. The estimated effect of the 90/75 ratio is in line with the findings in Rooth and Stenberg (2012), whereas the effect of the 50/10 ratio differs from their results. Results from Tables 2 and 3 thus indicate that inequality in different parts of income distribution have different effects on the income growth rate.

As such, our results corroborate the findings of Voitchovsky (2005). The results also suggest that a single measure of income inequality (such as the Gini coefficient) may be insufficient to capture the effects of income inequality on the income growth rate.

We also performed a sensitivity analysis to test the robustness of the results; our concern was that the crisis in Sweden during the early 1990s could potentially affect the estimated relationship between income inequality and income growth. Therefore, it is important to consider the downturn in the Swedish economy during 1990-1993. Hence, we checked the robustness of the results by excluding the data for 1992-1994, as they potentially can influence the estimated growth-inequality relationship, and re-estimated our models for the period 1995-2007 using a four-year lag instead of five-year lag[17]. The most important findings are summarised as follows[18]. When measuring income inequality solely by using the Gini coefficient or the different top income shares, the results are consistent with those presented in Tables 2 and A4. However, when the Gini coefficient, the 90/75 ratio, and the 50/10 ratio are examined simultaneously, we notice that the effect of the 50/10 ratio becomes positive but not significant. Therefore, the effect of the 90/75 ratio seems to be fairly stable, while the effect of the 50/10 ratio is more sensitive to the choice of time period for estimation. However, except for this difference, the results do not change much by comparison with those based on the original data. Therefore, we have not found any strong evidence suggesting that including the crisis years 1992 and 1993 leads to serious problems.

4. Conclusions

The main objective of the present paper is to investigate the effects of income inequality on the income growth rate using panel data from 283 Swedish municipalities during the years 1992-2007. We use several measures of inequality: the Gini coefficient, different top income shares, as well as measures of inequality in the lower (measured by the 50/10 ratio) and upper (measured by the 90/75 ratio) ends of the income distribution. We estimate a single equation panel data fixed-effects model, which relates the income growth rate to income inequality, and other control variables indicating local policy decisions, human capital, and socio-economic and demographic structures.

[17] We use a 4 year lag in order to have the same number of points of observations as our previous analysis; the 4 year lag also allows us to check the robustness of the results based on 5 year lags discussed above.

[18] The results of the sensitivity analysis are not reported in this study but they are available from the author upon request.

The main finding is that income inequality has a positive effect on the subsequent income growth at the municipal level in Sweden. The positive effect of the Gini coefficient supports previous findings based on Swedish panel data at the county level (Nahum, 2005). It also supports earlier results based on cross-country panel data. We also find that inequality measured as $Top25$, $Top15$, $Top10$, $Top5$ and $Top1$, respectively, has a positive effect on the income growth rate. A striking result is that the estimated interaction effect between the Gini coefficient and level of average income is positive and significant. As such, the larger the average income, the stronger will be the positive growth-effect of increased income inequality, ceteris paribus. A similar result was found on cross-country data by Barro (2000). Furthermore, and in line with Voitchovsky (2005), the results show that top end inequality measured by the 90/75 ratio has a positive effect on the income growth rate, while the bottom end inequality measured by the 50/10 ratio has a negative effect.

Several important questions remain to be addressed in future research. A natural next step of this empirical investigation would be to try to identify the different channels through which inequality in different parts of income distribution influences the growth process. In addition, an important limitation of the current analysis is the time span of the data that is used to estimate our model. It would be interesting to replicate our study for a longer time series to check more thoroughly if the estimated relationship between income inequality and income growth is sensitive to the choice of estimation period.

Appendices

Appendix A1: Gini coefficient

Consider a portion of a population of N individuals with a total income of A into a subclass of size N_i with respective income A_i, $i = 1,...,k$ (in our paper $k = 18$), we define:

$$p_i = \frac{1}{N}\sum_{j=1}^{i} N_j \qquad i = 1,...,k,$$

$$q_i = \frac{1}{A}\sum_{j=1}^{i} A_j \qquad i = 1,...,k,$$

with $p_0 = 0$ and $q_0 = 0$. Observe that p_i and q_i correspond to the cumulative share of population and income for class *i,* respectively.

We define

$$C = \frac{1}{2} - \frac{1}{2}\sum_{i=1}^{k}(q_i + q_{i-1})(p_i - p_{i-1})$$

Moreover, looking at the Figure A1, we see that the maximum value of the above expression is $C_{max} \cong \frac{1}{2}$. The Gini coefficient is defined as $G = \frac{C}{C_{\max}} \cong 1 - \sum_{i=1}^{k}(q_i + q_{i-1})(p_i - p_{i-1})$.

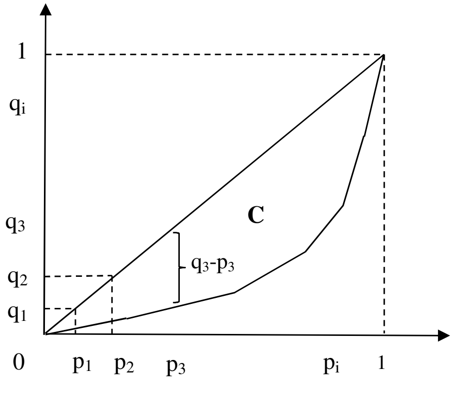

Figure A1: Graphical representation of the inequality distribution of income

Table A1: Descriptive statistics (1992-2007)

Variable		Mean	Std. Dev.	Min	Max
$y_{i,t}$	overall	1.021	0.016	0.962	1.092
	between		0.003	1.013	1.035
	within		0.016	0.960	1.078
$Y_{i,t}$	overall	65.840	10.778	45.895	151.325
	between		7.676	55.493	119.110
	within		7.578	37.945	98.056
$G_{i,t}$	overall	0.159	0.055	0.073	0.552
	between		0.045	0.100	0.464
	within		0.031	0.046	0.255
$Tax_{i,t}$	overall	31.349	1.325	25.700	34.410
	between		1.056	27.643	33.198
	within		0.803	28.615	34.406
$Exedu_{i,t}$	overall	0.368	0.037	0.224	0.561
	between		0.032	0.262	0.504
	within		0.018	0.258	0.445
$Exfam_{i,t}$	overall	0.065	0.027	0.011	0.215
	between		0.023	0.024	0.163
	within		0.013	0.020	0.183
$Exchil_{i,t}$	overall	0.148	0.039	0.059	0.326
	between		0.032	0.084	0.256
	within		0.021	0.089	0.256
$Exeld_{i,t}$	overall	0.357	0.069	0.089	0.532
	between		0.059	0.152	0.491
	within		0.036	0.204	0.443
$Exp_{i,t}$	overall	12.000	2.177	6.231	22.588
	between		1.084	9.296	15.788
	within		1.889	6.656	21.817
$Edu_{i,t}$	overall	0.061	0.033	0.020	0.261
	between		0.029	0.029	0.229
	within		0.016	0.010	0.129
$Herf_{i,t}$	overall	0.283	0.058	0.170	0.560
	between		0.049	0.193	0.489
	within		0.032	0.128	0.430
$Cons_{i,t}$	overall	0.467	0.123	0.107	0.889
	between		0.116	0.159	0.860
	within		0.042	0.336	0.652
$Age7\text{-}15_{i,t}$	overall	0.118	0.013	0.063	0.164
	between		0.010	0.070	0.146
	within		0.008	0.089	0.138
$Age65_{i,t}$	overall	0.190	0.038	0.059	0.302
	between		0.037	0.082	0.286
	within		0.008	0.152	0.229
$Dens_{i,t}$	overall	125.675	414.497	0.241	4228.241
	between		414.762	0.266	3950.624
	within		18.714	169.588	462.611

Note: *Overall* refers to the whole dataset. *Between* refers to the variation of the means to each municipality (across time periods). *Within* refers to the variation of the deviation from the respective mean to each municipality.
Observations: N=4528 for the *overall*, T=16 for *within* sample and n= 283 for *between.*
Between statistics are calculated on the basis of summary statistics of 283 municipalities, regardless of time period, while *Within* statistics are calculated on the basis of summary statistics of 16 time periods, regardless of municipalities (more details in note of Table A3). See Note in the next page the description of the *Overall, Within* and *Between* statistics.

Table A2: Descriptive statistics (1992-2007) top and bottom inequality income distribution

Variable		Mean	Std. Dev.	Min	Max
$Top1\%_{i,t}$	overall	0.020	0.016	0.011	0.026
	between		0.008	0.012	0.019
	within		0.014	0.013	0.017
$Top5\%_{i,t}$	overall	0.027	0.023	0.020	0.032
	between		0.010	0.014	0.029
	within		0.017	0.013	0.024
$Top10\%_{i,t}$	overall	0.150	0.035	0.130	0.226
	between		0.021	0.122	0.252
	within		0.013	0.005	0.025
$Top15\%_{i,t}$	overall	0.211	0.035	0.150	0.432
	between		0.021	0.172	0.252
	within		0.028	0.125	0.430
$Top25\%_{i,t}$	overall	0.304	0.028	0.250	0.432
	between		0.016	0.256	0.380
	within		0.022	0.222	0.428
$90/75_{i,t}$	overall	1.192	0.019	1.164	1.224
	between		0.012	1.174	1.221
	within		0.015	1.154	1.228
$50/10_{i,t}$	overall	3.221	0.600	1.639	3.997
	between		0.435	2.187	3.974
	within		0.414	1.628	4.459

Note: *Overall* refers to the whole dataset. The total variation (around grand mean $\bar{x} = 1/NT \sum_i \sum_t x_{it}$) can be decomposed into *within* variation over time for each individual country (around individual mean $\bar{x}_i = 1/NT \sum_t x_{it}$) and *between* variation across countries (for $\bar{x}$ around $\bar{x}_i$). The corresponding decomposition for the variance is

Within variance: $s_W^2 = \frac{1}{NT-1} \sum_i \sum_t (x_{it} - \bar{x}_i)^2 = \frac{1}{NT-1} \sum_i \sum_t (x_{it} - \bar{x}_i + \bar{x})^2$;

Between variance: $s_B^2 = \frac{1}{N-1} \sum_i (x_i - \bar{x})^2$;

Overall variance: $s_O^2 = \frac{1}{NT-1} \sum_i \sum_t (x_{it} - \bar{x})^2$.

The second expression for s_w^2 is equivalent to the first, because adding a constant does not change the variance, and it is used at times because $x_{it} - \bar{x}_i + \bar{x}$ is centered on $\bar{x}$, providing a sense of scale, whereas $x_{it} - \bar{x}_i$ is centered on zero.

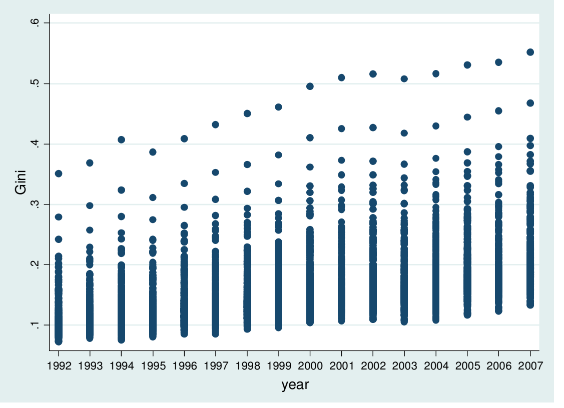

Figure A2: Distributions of Gini coefficients across municipalities 1992-2007

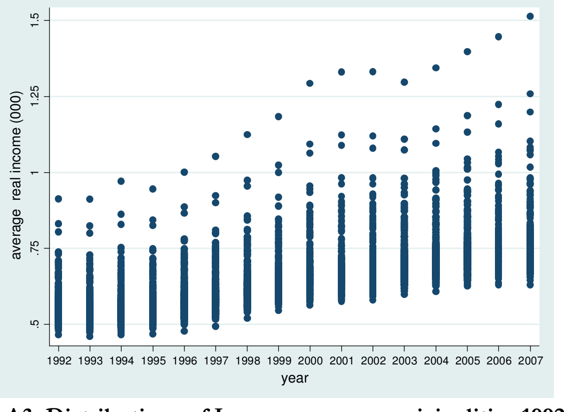

Figure A3: Distributions of Income across municipalities 1992-2007

Note: Each dot represents one municipality. For each year, the Gini coefficient (Fig. A2) and the real average income (Fig. A3) across municipalities is distributed between a minimum and a maximum value. From Fig. A2 we notice that the Gini coefficient increases over time across municipalities; the minimum value in 1992 was lower than the minimum value in 2007. Moreover, the difference between the lowest and the highest values of Gini coefficient has increased from 1992 to 2007. A similar trend is visible in the distribution of average income in Figure A3.

Table A3: Fixed Effects estimations top earners

	Top10		Top5		Top1	
	(i)	(ii)	(iii)	(iv)	(v)	(vi)
	ln(y)					
Constant	0.306*** (5.30)	0.378*** (6.53)	0.228*** (4.01)	0.300*** (5.26)	0.105 (1.95)	0.136* (2.55)
$G_{i,t-T}$	0.676*** (-10.78)	0.747*** (-11.92)	0.455*** (-8.66)	0.525*** (-9.94)	0.159*** (-6.65)	0.157*** (-6.61)
$Y_{i,t-T}$	0.257*** (-7.42)	0.680*** (-10.71)	0.252*** (-7.05)	0.668*** (-10.44)	0.321*** (-9.37)	0.657*** (-10.12)
$Tax_{i,t-T}$	-0.0516 (-0.78)	-0.0814 (-1.25)	-0.0545 (-0.82)	-0.088 (-1.34)	-0.036 (-0.54)	-0.052 (-0.79)
$Exedu_{i,t-T}$	0.242*** (4.26)	0.228*** (4.05)	0.214*** (3.75)	0.201*** (3.55)	0.164** (2.88)	0.143* (2.53)
$Exfam_{it-T}$	0.650*** (10.58)	0.633*** (10.41)	0.636*** (10.29)	0.620*** (10.13)	0.596*** (9.62)	0.577*** (9.37)
$Exchil_{i,t-T}$	0.371*** (5.90)	0.326*** (5.22)	0.362*** (5.72)	0.316*** (5.03)	0.335*** (5.26)	0.300*** (4.72)
$Exeld_{i,t-T}$	0.391*** (7.55)	0.374*** (7.28)	0.357*** (6.87)	0.339*** (6.60)	0.307*** (5.95)	0.285*** (5.53)
$Exp_{i,t-T}$	0.085*** (-13.80)	0.092*** (-14.89)	0.092*** (-14.77)	0.098*** (-15.77)	0.109*** (-19.63)	0.117*** (-20.62)
$Edu_{i,tT}$	2.274*** (21.37)	2.161*** (20.35)	2.248*** (20.57)	2.150*** (19.75)	2.055*** (19.65)	1.933*** (18.27)
$Herf_{it-T}$	0.101*** (5.64)	0.114*** (6.43)	0.103*** (5.73)	0.116*** (6.50)	0.115*** (6.35)	0.126*** (6.98)
$Cons_{i,t-T}$	-0.326*** (-24.62)	-0.325*** (-24.81)	-0.325*** (-24.33)	-0.324*** (-24.52)	-0.320*** (-23.91)	-0.319*** (-23.94)
$Age65_{i,t-T}$	0.042 (0.45)	0.103 (1.10)	-0.023 (-0.24)	0.033 (0.35)	0.002 (0.02)	0.040 (0.42)
$Dens_{i,t-T}$	-0.125*** (-3.66)	-0.112** (-3.29)	-0.157*** (-4.58)	-0.145*** (-4.29)	-0.148*** (-4.27)	-0.143*** (-4.15)
$(G{\cdot}Y)_{i,t-T}$		0.922*** (7.92)		0.919*** (7.80)		0.713*** (6.08)
N of observation	3113	3113	3113	3113	3113	3113

Note: t-values in parentheses. Significance level: "*"p<0.05, "**", p<0.01, "***" p<0.001

Table A4: Estimations for Poor and Rich and municipalities (T=5)

	Split sample		Whole sample	
	FE		FE	
	Rich	Poor	Poor	Dummy
Constant	-2.800**	-0.408***		
	(-3.65)	(-5.17)		
$G_{i,t-T}$	1.181*	0.736***	1.123***	0.203***
	(2.33)	(5.56)	(7.80)	(2.79)
$Y_{i,t-T}$	-0.929***	-0.491***	-0.534***	-0.162**
	(-5.03)	(-13.83)	(-7.44)	(-3.05)
$Tax_{i,t-T}$	-0.158	-0.068	-2.186	-0.135
	(-0.25)	(-0.78)	(-1.38)	(-1.30)
$Exedu_{i,t-T}$	2.187**	0.140*	0. 204	0.103
	(3.03)	(2.07)	(3.26)	(2.38)
$Exfam_{it-T}$	2.385**	0.105	0. 421**	0. 347**
	(3.21)	(1.45)	(2.86)	(2.60)
$Exchil_{i,t-T}$	2.176**	0.183*	0.150**	0. 136**
	(3.20)	(2.29)	(2.31)	(2.55)
$Exeld_{i,t-T}$	1.862**	0.127*	0.161*	0. 156*
	(2.94)	(2.02)	(2.57)	(2.02)
$Exp_{i,t-T}$	0.155*	-0.013	0.382***	0.707***
	(2.23)	(-1.29)	(2.70)	(2.64)
$Edu_{i,tT}$	2.598	1.862***	0.537***	0.401***
	(2.09)	(8.05)	(2.62)	(2.17)
$Herf_{i,t-T}$	0.041	-0.024	0.011	0.036
	(0.33)	(-0.97)	(0.03)	(0.79)
$Cons_{i,t-T}$	-0.030	-0.169***	-0.191***	0.058**
	(-0.29)	(-9.67)	(-7.60)	(2.92)
$Age65_{i,t-T}$	0.253	0.310	0.627***	0.175*
	(0.43)	(1.81)	(4.68)	(2.09)
$Dens_{i,t-T}$	0.064	-0.421**	-0.205	0.026**
	(0.30)	(-3.24)	(-1.95)	(2.87)
N of observations	340	823	1131	

Note: t-values in parentheses. Significance level: "*"p<0.05, "**", p<0.01, "***" p<0.001

In the model for the "whole sample", the parameter estimates for the rich municipalities can be calculated by summing the estimates in the "Poor" and "Dummy" columns. Also, to save space, the fixed effects are not presented.

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