CONVERGENCE IN PER CAPITA INCOME AND MIGRATION ACROSS THE SWEDISH COUNTIES 1906-1990

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Abstract

This paper finds strong and robust evidence of convergence in per capita income across the twenty-four Swedish counties 1906-1990. It is found that migration has a positive effect, albeit small, on the speed of convergence. Holding net migration constant, the estimated speed of convergence is around 3 percent per year, which is higher than estimates obtained by Barro and Sala-i-Martin (1991,1992) for other regional data sets. One likely explanation of this finding is that the current study, as opposed to previous studies, adjusts incomes to account for regional differences in cost of living.

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1. INTRODUCTION

One of the most debated issues in the economic growth literature during the last ten years is whether per capita incomes in different countries are converging. In their seminal papers both Romer (1986) and Lucas (1988) cited the lack of observed cross-country convergence as evidence against the neoclassical model and in favor of their theories of endogenous growth. Since then numerous cross-country studies (e.g. Barro (1991) and Mankiw, Romer and Weil (1992)) have found a negative relation between

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initial per capita income and subsequent growth in per capita income after having controlled for other variables that affect per capita income growth. In other words, these studies find conditional b-convergence, which is the prediction of the neoclassical model. Moreover, Barro and Sala-i-Martin (1991,1992,1994) find unconditional b-convergence across regions in the U.S., Western Europe, and Japan, i.e. they find that poorer regions tend to grow faster than rich ones in per capita terms without the inclusion of conditioning variables. Barro and Sala-i-Martin argue that the differences with respect to variables that determine economies' steady states, i.e. with respect to preferences, technology and institutions, are likely to be smaller for regions within countries than across countries. This relative homogeneity implies that unconditional convergence is more probable across regions within countries than across countries.

b-convergence is consistent not only with the neoclassical growth model but also with some endogenous growth models. For one-sector growth models the key to b-convergence is the existence of diminishing returns to reproducible inputs whereas the key to endogenous growth is the violation of the Inada condition. There are one-sector models that generate both endogenous steady-state growth and b-convergence (see e.g. Barro and Sala-i-Martin (1994, ch. 1)). b-convergence is however not consistent with the one-sector AK-models (see e.g. Sala-i-Martin (1994)), which may explain why there has been so much interest in the existence of b-convergence across economies. Moreover, Mulligan and Sala-i-Martin (1993) show that two-sector models, such as the Lucas-Uzawa model, could fit the regression statistics on b-convergence given that the initial income of an economy is correlated with the degree of imbalance among the sectors.

Another class of growth models that is often put forward as an explanation for b-convergence is models of technological diffusion. The main idea here is that economies tend to catch up with the technological leader because technological progress is more rapid in the follower economies than in the leading economy since imitation and implementation of discoveries are cheaper than innovation (see e.g.

Barro and Sala-i-Martin (1994, ch. 8). Moreover, diminishing returns and technological diffusion are not mutually exclusive sources of convergence.

The main purpose of this paper is to test for b-convergence in per capita income across the twenty-four Swedish counties 1906-1990. In addition to testing for unconditional b-convergence, I control for interregional migration as well as for shocks to the agricultural sector. I also discuss the potential effects on b-convergence of some institutional features of Sweden; namely, of the solidarity wage policy of the unions and of the regionally redistributive policies (or lack thereof) of the central government. Moreover, like most of the empirical literature on convergence, this paper does not attempt to discriminate between different growth models that predict b-convergence.

Another purpose of the paper is to study s-convergence in per capita income across the Swedish counties. The concept of s-convergence deals with how the distribution of income or product across economies evolve over time whereas b-convergence deals with the mobility of income within the same distribution. b-convergence is a necessary but not a sufficient condition for s-convergence (see e.g. Sala-i-Martin (1994)).

The Swedish counties provide strong and robust evidence of convergence. Controlling for migration and shocks to the agricultural sector does not change the results on b-convergence in any major way. It is found that interregional migration has a positive effect, albeit small, on the speed of convergence. These results are not favorable for the one-sector AK-model. Holding net migration constant, the estimate of the speed of convergence is around 3 percent per year, which is higher than estimates obtained by Barro and Sala-i-Martin (1991,1992) for other regional data sets. One likely explanation of this finding is that the current study, as opposed to previous studies, adjust incomes for regional differences in cost of living.

The rest of the paper is organized in the following way: Sections 2 discusses the basic regression equation. Section 3 describes the data set. Section 4 presents the empirical results. Section 5 concludes.

2. THEORETICAL BACKGROUND

Even though my purpose is not to test a specific growth model or class of growth models, I discuss briefly in this section the basic regression equation, that is used for testing for b-convergence, on the basis of the neoclassical framework.

Log-linearizing the differential equations of a standard closed economy Ramsey model with labor-augmenting technological progress (see e.g. Barro and Sala-i-Martin (1994, ch. 2)) around steady state yields the following equation,

$$\log \hat{y}(t) = e^{-bt} * \log \hat{y}(0) + (1 - e^{-bt}) * \log \hat{y}_{ss}, \qquad b > 0 \tag{1}$$

where $\hat{y}$ is output or income per effective worker, $\hat{y}_{ss}$ is the steady-state level, and b is the rate of convergence, which is determined by the parameters of technology and preferences.

Rewriting some of the variables expressed in units of effective labor in equation (1) into per capita units, setting $t = T$ and rearranging, yields an expression for the average growth rate of per capita output or income,

$$(1/T) * \log[y(T) / y(0)] = x + \left[(1 - e^{-bT}) / T\right] * \log[\hat{y}_{ss} / \hat{y}(0)] \tag{2}$$

where x is the exogenous rate of labor-augmenting technological progress. Given x and $\hat{y}_{ss}$, the average per capita income growth rate is inversely related to initial per capita income, i.e. there is conditional b-convergence.

The assumption of a closed economy model may appear hard to justify when dealing with regions within a country. At the other extreme, an open economy Ramsey model with perfect capital mobility generates a number of counterfactual conclusions, one of which is infinite speeds of convergence in output (see e.g. Barro and Sala-i-Martin (1994, ch. 3)). Barro, Mankiw and Sala-i-Martin (1992) attempt to reconcile the empirical evidence on convergence with an open economy neoclassical model.

The important assumption is that only a fraction of the capital stock, which consists of physical and human capital, can be used as collateral on interregional and international loans. The main message of the model is that the dynamic properties are similar to those of the closed economy Ramsey model: A credit-constrained economy experiences conditional convergence in both output and income and the speed of convergence is only somewhat higher than in the closed economy model given that at most about half of the capital stock can serve as collateral on foreign loans. A problem is however that the capital stock and output of an unconstrained economy instantaneously adjust to their steady-state levels. If adjustment costs to investment were introduced this would remove the infinite speeds of convergence[1]. Since the closed economy neoclassical model does not appear to generate too implausible predictions, I will follow Barro and Sala-i-Martin (1991,1992) and use it as a starting point for the analysis.

To derive the statistical model, consider a discrete period version of equation (2) that applies to economy i and is augmented by a disturbance term ,

$$\log[y_{i,t} / y_{i,t-1}] = a - (1 - e^{-b}) * [\log y_{i,t-1} - x * (t-1)] + u_{i,t} \tag{3}$$

where $u_{i,t}$ is the disturbance term. I make the assumption that the economies are equal with respect to technology and preferences, which implies that b is the same across economies and $a = x + (1 - e^{-b}) * \log \hat{y}_{ss}$.[2]

The statistical model that is used for testing for b-convergence is given by equation (4). It is implied by equation (2) and (3). The average growth rate for economy i between two points in time, t_0 and $t_0 + T$, is given by,

$$(1/T) * \log[y_{i,t_0+T} / y_{i,t_0}] = c - \left[(1 - e^{-bT}) / T\right] * \log y_{i,t_0} + u_{i,t_0,t_0+T} \tag{4}$$

[1] For an in-depth discussion on neoclassical models that include open economy considerations and adjustment costs to investment see Barro and Sala-i-Martin (1994, ch. 3).
[2] It can be noted that the level of technology does not influence the value of b in the Ramsey model.

where $u_{i,t_{0},t_{0}+T}$ is the error term and $c = x + [(1 - e^{-bT}) / T] * [\log \hat{y}_{ss} + x * t_0]$. The intercept is increasing in t_0 due to technological progress.

Barro and Sala-i-Martin's (1991,1992) find convergence for different sets of regions using data on per capita income and per capita output at a rate of about 2 percent per year. The estimated rates of convergence for per capita income and per capita output are similar. As is well known, a rate of convergence of 2 percent per year seems to be inconsistent with a closed economy neoclassical model with a conventional capital share of around 0.3. Given a set of reasonable parameter values, a closed economy Ramsey model with a capital share of 0.3 generates a speed of convergence of about 6 percent per year (Barro and Sala-i-Martin, ch. 2)). If, in addition, one considers mobility of capital and labor across economies as well as technological diffusion from rich to poor economies, one would in general ,at least for per capita products, expect even higher speeds of convergence (see e.g. Barro et al. (1992a)).

3. DATA

Data on real per capita income, $y_{i,t}$, for the Swedish counties, $i = 1,\ldots,24$, for the period 1906-1990 is used. There are some gaps in the time series. I have annual observations for 1906, 1911-1912, 1916 and 1919-1990. The measure of income is gross per capita income net of government transfers[3]. There are no available county-specific price indices. To get a more accurate representation of each county's per capita income I adjust however the counties' real incomes, deflated by the national consumer price index, to account for differences in price levels across counties. Only

[3] Up to 1916 the measure of income is however taxable income net of transfers. Source: Royal Ministry of Finance, the Tax Assessments, various issues. The income concept 1919-1990 is gross income net of transfers. Source: Statistics Sweden, the Tax Assessments and the Income Statistics, various issues. To get per capita income I use data on the counties' population from Statistics Sweden, the Population Statistics, various issues. I have compared my data set with national GDP (based on Hassler, Lundvik, Persson and Söderlind (1992) and Statistics Sweden, National Accounts). The annual rate of change of the sum of the counties' incomes in real terms corresponds closely to the annual rate of change of real national GDP for the period 1919-1990.

differences in housing prices are accounted for. I have data on regional prices for food and fuel, and housing. The regional price differences for food and fuel are small, even at the beginning of the sample period (Ohlin (1924), Björklund and Stenlund (1984,1990), and Skedinger (1991)). As a result, no adjustment is made for differences in food and fuel prices. Housing costs do however differ across counties. As a proxy for the differences in housing costs I use data on rents (including heating costs) of a standard one room apartment in the county capitals for the period 1911-1946[4]. The time series on rents ends in 1946. For the period after 1946, I use data on the average sales prices of a one-family house in the counties. This data series starts only in the 1950s. The first observation is an average for the period 1952-1956. Annual data are available starting in 1957[5]. Data limitations force me to use two different non-overlapping series as proxies for housing costs. This potentially introduces measurement errors. In appendix B I discuss this issue. The conclusion is that the use of two different series is not likely to introduce any serious measurement errors. To adjust the counties' real incomes I create a vector, z_t. Housing costs are about twenty percent of the consumer's budget throughout the whole sample period for all counties[6]. I use this figure when calculating z_t ,

$$z_t = [(p_t / \bar{p}_t) - 1] * 0.2 + 1 \tag{5}$$

where p_t is a vector of sales prices of the one-family house or a vector of rents of the one room apartment in the counties at year t , $\bar{p}_t$ is the average value across counties. For example, for a county i , whose housing costs are twice that of the average value across counties, $z_{t,i}$ is 1.2. To adjust incomes, the logarithm of z_t is subtracted from

[4] The data on rents are not available annually. I have data on average rents for the periods 1912-1915, 1924-1925, 1928-1929, 1933-1934, 1939-1940 and 1945-1946. Source: Statistics Sweden, Census of Rents, various issues. For years that data on rents are not available I use the rents of the most adjacent period to adjust incomes, see appendix B.
[5] Data on housing prices are from Statistics Sweden, Statistical Journal, SM R and SM P, various issues.
[6] Statistics Sweden (1932, 1961), Consumer Prices and Index Computations 1931-1930 and 1931-1959 and Statistics Sweden, Statistical Yearbook, various issues.

the logarithm of a vector of the counties' real incomes deflated by the national consumer price index. When I refer to real incomes in the subsequent text I mean real incomes adjusted for differences in housing costs if not anything else is said.

Incomes for the period 1985-1990 are all adjusted on the basis of the relative housing prices in 1985. During the latter part of the 1980s housing prices ballooned partly because of a reformed tax system and a deregulated financial market. This price increase was not symmetric across counties. Housing prices increased relatively more in rich counties. This relative price change was however temporary in nature. In 1993 the relative housing prices across counties were about the same as they were in 1985. Since temporarily inflated housing prices are not good proxies for the cost of living for a majority of the citizens, which I attempt to adjust for, I have chosen to use the relative housing prices in 1985 as a basis for such an adjustment for the period 1985-1990 (For further discussion see footnote 7 in section 4).

4. EMPIRICAL RESULTS

The relative differences in per capita incomes across counties were large in the beginning of the century. In 1906 the per capita income, in 1980 year prices, was around 7.300 kronor in the richest county, whereas it was only about 800 kronor in the poorest county. More recent nominal estimates should however be more reliable. In 1919 the real per capita income was around 9.100 kronor in the richest county and about 2.200 in the poorest county. In 1990 the relative income differentials had narrowed dramatically. The real per capita income in 1990 was about 46.300 kronor in the richest county, and approximately 38.500 in the poorest county. The absolute difference between the richest and poorest county was about the same, around 7.000 kronor, in 1990 as it was in the beginning of the sample period.

Figure 1 and 2 provide strong evidence of convergence across the Swedish counties. Figure 1 shows a remarkably strong negative relation between the counties' average annual real per capita income growth rates 1906-1990 and the logarithm of real per capita incomes in 1906. Figure 2 plots the logarithm of the counties' real per capita incomes 1906-1990. In the upper window of figure 2 real incomes are adjusted for regional differences in housing costs whereas real incomes plotted in the lower window are not adjusted for such differences. Taking into account the regional differences in housing costs compresses the distribution of income. This happens because the richest counties have the highest housing costs. The correlation between the counties' real incomes (not adjusted for differences in housing costs) and relative housing costs, z_t, is high throughout the sample period. For example, the correlation between incomes in 1911 and relative average housing costs for the period 1912-1915 is 0.82 and the correlation between incomes in 1985 and relative housing costs for the same year is 0.88.

The statistical model in equation (4) is estimated by nonlinear least square. Table 1 shows the results. For the longest sample period, 1906-1990, the estimate of b is 0.041 (8.1). The t-value is given within parentheses.

As is well known, a measurement error in the initial level of real per capita income tends to bias the estimate of b upwards (De Long (1988)). One check on the importance of measurement errors can be performed by examining convergence starting at some later point in time for which, for example, nominal income estimates are more reliable. Table 1 shows that the estimates of b are similar for the three sample periods; 1906-1990, 1911-1990 and 1919-1990. Judging from this evidence the measurement errors are limited.

The overall sample is also divided up into subperiods which can be seen as a test of robustness. Rows 4-11 of table 1 show estimates for eight subperiods of the overall sample: 1906-1916, 1919-1930, 1930-1940, 1940-1950, 1950-1960, 1960-1970, 1970-1980 and 1980-1990. For all subperiods, except for 1980-1990, the estimate of b is significantly positive. The estimation method for the subperiods is nonlinear seemingly unrelated regression (SUR). Using SUR is a way to account for e.g. long-lasting sectoral shocks, since it allows for correlation of disturbances over time. It turns out however that this correlation is small which means that the SUR-estimates are similar to the ones obtained by applying nonlinear least squares to the equations individually.

If the eight subperiods are restricted to have the same b but individual constants, then the joint estimate of b is 0.037 (26.9). A likelihood ratio test however strongly rejects the hypothesis that b is the same for the subperiods. The estimated b is highest for the following subperiods: 1940-1950, 1960-1970, and 1970-1980. The instability of b across subperiods will be analyzed later in this section.

Equation (4) is also estimated when incomes are not adjusted for regional differences in housing costs. The results are similar to the ones obtained when incomes are adjusted for differences in housing costs in the sense that the estimates of b for the whole sample period as well for all but the last subperiod are significantly positive. A difference is however that when non-adjusted incomes are used the estimates of b are generally lower. For example, for the longest sample period 1906-1990 the estimate of b is only 0.027 (15.0). When incomes are not adjusted the estimate of b tends to decrease because the richest counties have the highest housing costs. Graphically this means that the fitted curve to the observations in figure 1, tilts: the intercept and the absolute value of the slope decrease. A flatter slope implies a slower speed of convergence. Moreover, if housing costs in poor counties relative to housing costs in rich counties fall over time, this would also contribute to a lower estimate of b if non-adjusted incomes are used. The housing costs of poor relative to rich counties did however not diverge during the whole sample period[7]

[7] The correlation between the growth rate of the counties' relative housing costs between 1912-1915 and 1985, $(\log z_{85} - \log z_{12-15})$, and the logarithm of real incomes in 1911 is only 0.03. The correlation between the growth rate of the counties' relative housing costs between 1912-1915 and 1990, $(\log z_{90} - \log z_{12-15})$, and the logarithm of real incomes in 1911 is 0.48, which indicates a fair amount of price divergence. This means that the relative housing prices in rich counties increased during the latter part of the 1980s. This was however a temporary relative price hike. This is shown by the fact that the correlation between the growth rate of the counties' relative housing costs between 1912-1915 and 1993, $(\log z_{93} - \log z_{12-15})$, and the logarithm of real incomes in 1911 is low, 0.16. For the subperiod, 1930-1940, that the estimate of b is higher when non-adjusted incomes are used is a period of price convergence, i.e. there is a negative correlation between the growth rate of the counties' relative housing costs and initial incomes.

That the estimates of b for the Swedish counties in general are higher when differences in cost of living are adjusted for is consistent with findings across countries. Barro and Sala-i-Martin (1994, ch. 12) test for conditional convergence across a large sample countries 1965-1985[8]. The GDP-figures are based on Summers and Heston (1993). The reported speed of convergence is about 3 percent per year[9]. Barro and Sala-i-Martin also replaces the GDP-figures from the Summers and Heston data set with the World Bank figures on GDP. In all other respects, i.e. with respect to countries included, sample period, and other explanatory variables, the estimation is the same. When the World Bank figures are used the estimate of the speed of convergence is only about half of the estimate obtained by using the Summers and Heston data set. As opposed to the World Bank, Summers and Heston attempt to adjust for cross-country differences in the cost of living by using observed prices of goods and services. The Summers and Heston data set should therefore be a more accurate description of each country's per capita output. Barro and Sala-i-Martin also find that the standard deviation of the logarithm of the countries' GDPs based on the Summers and Heston data set is lower than for the World Bank data set. One likely explanation of these findings is that poor countries tend to have relatively low prices for nontraded goods. These results parallel those of mine; when I adjust for regional differences in cost of living the estimated speed of convergence increases and the standard deviation of the logarithm of real per capita income falls (see figure 3).

Barro and Sala-i-Martin (1991,1992,1994) do not adjust incomes or products for regional differences in cost of living when they study convergence across regions. The results discussed above imply that one would expect Barro and Sala-i-Martin's regional estimates of b, that are around 2 percent per year, to be biased downwards. These estimates of b are also lower than the estimates for the Swedish counties.

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[8] The sample period is divided up into two ten-year subperiods; 1965-1975 and 1975-1985. 79 countries are included in the first subperiod and 92 countries in the second subperiod.

[9] Mankiw et al. (1992) estimate the speed of convergence between 1960 and 1985 for a sample of 75 countries using the Summers and Heston (1988) data set to be almost 0.02. They use a much smaller set of variables to account for differences in steady states than Barro and Sala-i-Martin (1994, ch. 12) do, which may be one reason why these studies yield different estimates.

4.1. Sectoral Shocks and Convergence

The unstable pattern of $\hat{b}$ in table 1 could reflect sectoral shocks. Shocks that alter the sectoral composition of the counties' incomes among sectors of different average levels of productivity affect the estimated speed of convergence if the induced sectoral changes are correlated with initial levels of per capita income. For example, shocks to agriculture, which traditionally has been a sector of low productivity[10], are more important for counties dominated by agriculture. These counties tend also to be the poorest ones. The correlation between incomes in 1911 and the share of the population (or labor force)[11] engaged in agriculture in 1910 is -0.94. This correlation falls monotonically throughout the sample period. In 1980 it is -0.72. As a result, the estimate of b captures the effect on income of shocks to agriculture. Moreover, we also tend to have a problem of serial correlation. Only shocks to the agricultural sector are considered here. To account for such shocks that thus far have been included in the error term, I include the proportion of the population (or labor force) engaged in agriculture at the beginning of the sample period in the basic regression model in equation (4).

The estimated bs for the whole sample period when initial agricultural share is included in the regression model are close to the ones obtained in the basic regression model. The estimated coefficients of the agriculture share variable are about zero and insignificant. Table 2 displays the results. Moreover, the correlation between the share of the population in agriculture in 1910 (1920) and the absolute value of the change in the share of the population in agriculture between 1910 (1920) and 1990 is in excess of 0.99. If we use the change in the agricultural share as a measure of structural

[10] There seems to be a high productivity differential between the agricultural sector and the non-agricultural sector, even though it is declining over time. I measure agricultural productivity as agricultural national income divided by population in agriculture and non-agricultural productivity as non-agricultural national income divided by population in non-agricultural activities. The ratio of non-agricultural to agricultural productivity was 3.0 in 1920, 3.5 in 1930, 2.4 in 1940, and 1.5 in 1950. The data is from the Population Censuses and the Income Statistics, various issues.

[11] For the years 1910, 1920, 1930 and 1940 I use the share of the population in agriculture. For 1950, 1960, 1970, 1980 and 1990 I use the share of labor force in agriculture. Source: Statistics Sweden, the Population Censuses and the Income Statistics, various issues.

change, this means that the estimates of b for the whole sample period in table 2 are net of the structural effects on income from shifts of labor from agriculture towards sectors of higher productivity[12]. For many subperiods, particularly in the beginning of the sample period, initial agricultural share and the change in agricultural share are however not highly correlated.

As one would expect that controlling for shocks to agriculture would make b more stable across the subperiods, I test the hypothesis that the speed of convergence is the same across subperiods. In the restricted model the eight subperiods are restricted to have the same b but individual constants and individual coefficients of the agriculture share variable. The joint estimate of b is 0.046 (9.6), which is higher than the joint estimate in the basic regression model[13]. The hypothesis of equal speeds of convergence for the subperiods is however still rejected by a likelihood ratio test at the 5 percent level. (The p-value is 0.012).

For some subperiods the estimated coefficient of the agriculture share variable is significantly negative, which means that agricultural counties on average had lower per capita income growth, holding initial per capita income constant.

To conclude, I find b-convergence net of the effects on income from shifts of labor from agriculture towards sectors of higher productivity for the whole sample period. The estimated bs for the whole sample period are about the same as in the basic regression model. For the subperiods, the estimates of b tend however to be higher. Moreover, the hypothesis of equal bs across subperiods is still rejected in the augmented regression model. By and large, these results are similar to those found by

[12] Structural change from agriculture towards manufacturing and service industries is of course an important feature of the sample period. In 1910, 48.4 percent of the population were engaged in agricultural activities. In 1990 this share had dropped to 1.4 percent. In 1910 the highest share of the population in agriculture among the counties was 76 percent, whereas the lowest was 18 percent. In 1990 the highest share was 7 percent and the lowest was 0.4 percent.

[13] In table 1 the first subperiod is 1906-1916. Since I only have data on the share of population in agriculture for every ten years, i.e. for 1910, 1920, etc., I have replaced the subperiod 1906-1916 with 1911-1919 in table 2. The estimate of the speed of convergence for the subperiod 1911-1919 is insignificant regardless of whether initial agricultural share is included or not. For two subperiods, 1919-1930 and 1930-1940, the estimated speed of convergence turn insignificant at the 5 percent level in the augmented regression model. The presence of multicollinearity might explain this. Multicollinearity is indicated by R square values of 48.7 and 39.3 percent respectively and by insignificant coefficients of initial income and initial agricultural share.

Barro and Sala-i-Martin (1992a) for the U.S. States with the exception that Barro et al. did not reject the hypothesis of equal bs across subperiods when sectoral shocks were controlled for.

4.2. s -convergence

Figure 3 plots the standard deviation of the logarithm of per capita income [14]. For the Swedish counties, the standard deviation falls from 0.45 in 1906 to 0.03 in 1990. Hence, there is s -convergence. The fall is however not monotonic over time. The dispersion increases during the 1920s. A similar pattern is found for the U.S. States. Barro and Sala-i-Martin (1992a) explain this with rapidly falling agricultural prices that adversely affected the relatively poor agricultural states. Moreover, the dispersion increases somewhat during the 1980s, even though the rise in the dispersion is reversed at the very end of the decade. This increase is not unique for Sweden. Also the U.S., Japan and some of the West European countries in the sample of Barro and Sala-i-Martin (1994, ch.11) experience increased dispersion in per capita income or product across regions during the 1980s.

4.3. Alternative Explanations of the Results

In this section I discuss some alternative explanations, other than diminishing returns to capital and technological diffusion, to the results on convergence. I focus on trade union behavior, central government policies and migration.

[14] If the rate of growth is constant across counties that start from different levels, this measure of dispersion stays constant.

4.3.1. Solidarity Wage Policy

One institutional feature of the Swedish Labor Market has been the solidarity wage policy of the unions. The solidarity wage policy together with centralized wage bargaining, which is a prerequisite for its implementation, started being introduced in the 1960s. According to Hibbs (1990), the objective of the solidarity wage policy up to the mid-60s was "equal pay for equal work", i.e. a leveling of wages among jobs of comparable difficulty, risk and skill. In the mid-60s the objective shifted towards more general wage equalization, i.e. special attention was given to low-income workers. The wage dispersion within industries as well as between industries fell dramatically between the mid-60s and the end of the 1970s. This is often attributed to the solidarity wage policy. Hibbs however points out that the wage dispersion could have fallen as a result of conventional supply and demand forces. The centralized bargaining was gradually eroded during the 1980s and the central union organizations began to favor "fair pay differentials" (Hibbs (1990)). The wage dispersion within industries as well as between industries increased also during the 1980s.

In table 1 the estimates of b for the subperiods 1960-1970 and 1970-1980 are higher than the estimates of b for other subperiods. An interesting question is therefore whether the speed of convergence in per capita income during the era of solidarity wage policy is *significantly* higher than the speed of convergence for other periods. The null hypothesis is that the speeds of convergence during the 1960s and 1970s do not differ from the speeds of convergence for other periods. This hypothesis is tested by a likelihood ratio test. In the unrestricted model b is the same for all subperiods, except for 1960-1970 and 1970-1980. These two subperiods share a common b. In the restricted model b is the same for all subperiods. The likelihood ratio statistic is 9.90, which is well above the 5 percent critical value of 3.84 of the c^2 distribution with one degree of freedom. (The p-value is 0.0016). Hence, I reject the null hypothesis of equal bs. I perform the same test when initial agricultural share is included in the income growth regressions. The likelihood statistic is 4.91 which implies that the hypothesis of equal bs again is rejected at the 5 percent level (the p-value is 0.027).

The joint estimates for the subperiods, when the 1960s and 1970s are excluded from the two samples, are 0.034 (15.6) and 0.040 (6.8) respectively. Thus, even when I exclude periods during which centralized trade union activity might have influenced the regional income distribution, the estimated speeds of convergence are still higher than estimates of other regional convergence studies.

4.3.2. National Government Policies

Redistribution from rich to poor counties through central government policies could conceivably have contributed to the observed convergence. Due to lack of relevant and accessible data, I will however not be able to address this issue in any depth. I will here only refer to a government study (SOU 1989:65) that allocated *all* expenditures of the central government and the social security funds to the counties for one fiscal year, 1985-1986. Expenditures in, for example, the defense budget for a particular regiment were allocated to the county in which the regiment was situated. There was no relation between per capita expenditures across counties and the counties' per capita incomes (net of transfers). The estimated correlation coefficient was -0.01. Hence, central government expenditures were not redistributive across counties for this particular fiscal year.

4.4. Migration and Convergence

Migration of labor with little human capital from poor regions, where capital-labor ratios and wages are low, to richer regions, where capital-labor ratios and wages are high, tends to increase the capital-labor ratios and wages in the regions of departure and to decrease the capital-labor ratios and wages in the regions of destination. Thereby, migration contributes to convergence in per capita income. In this section I discuss how much of the observed convergence can be explained by migration flows. As a theoretical framework for the analysis an overlapping-generation neoclassical model with migration is used. Absence of altruism between existing residents and immigrants motivates the use of an overlapping-generation model. The model is similar to that of Blanchard (1985). The model is from Barro and Sala-i-Martin (1994, ch. 9). It is briefly described in appendix A. The model is closed from international capital flows other than the human and physical capital migrants carry with them. Migrants are not allowed to maintain financial claims on foreign-source income. Moreover, given that the Swedish counties share the same steady state it is plausible to assume that the steady state of the overlapping-generation model is identical to that of a standard closed economy Ramsey model without migration. (For details, see appendix A). The speed of convergence towards steady state tends however to be higher in the model with migration. The effect of migration on the speed of convergence depends positively on the sensitivity of net migration to income differentials as well as negatively on the amount of capital per migrant, that migrants carry with them, relative to the capital-labor ratio in the domestic economy. A parameter, b, which is positively related to the speed of convergence, captures this. b is given by,

$$b = a * [1 - (k \ / \ k)] * \P m(\hat{y}) / \P \log \hat{y} \qquad (6)$$

a is the capital share and k / k is the amount of capital per migrant relative to the per capita capital in the domestic economy. $m(\hat{y})$ is the postulated net migration function. If k / k is one or if net migration is not affected by income differentials across economies, i.e. if $\P m(\hat{y}) / \P \log \hat{y} = 0$, then b is zero and the speed of convergence is no different from the speed of convergence in the standard closed economy Ramsey model. If, on the other hand, net migration is very sensitive to income differentials and $k / k < 1$, then the speed of convergence is strongly influenced by migration. To get a value of the sensitivity of net migration to per capita income differentials across the Swedish counties, the following statistical model is estimated,

$$m_{i,t_0,t_0+T} = c + d * \log y_{i,t_0} + v_{i,t_0,t_0+T} \tag{7}$$

where m_{i,t_0,t_0+T} is the average annual net migration rate for county i between time t_0 and $t_0 + T$. The rate is calculated as the share of net migration to population. If $m_{i,t_0,t_0+T} > 0$, then immigration is larger than emigration for county i [15]. v_{i,t_0,t_0+T} is the error term.

There is a positive relationship between initial real per capita income and subsequent net migration for the whole sample period 1911-1990. The estimate of d is 0.007 (6.3). Table 3 displays the results. The coefficient d measures the percentage change in population, through net migration, of one percent change in per capita income, holding constant the effect of income on fertility and mortality. The estimate of d for the whole sample period, implies therefore, ceteris paribus, that one percent increase in a county's per capita income raises net migration only by enough to increase the county's annual rate of population growth by 0.007 percentage points.

For all but the last two subperiods, 1970-1980 and 1980-1990, the estimates of d are significantly positive. For the sample period 1970-1980 the estimate of d is significantly negative. The estimate of d for the last subperiod, 1980-1990, is also negative but insignificant[16]. Allowing for individual constants for the subperiods, a likelihood ratio test strongly rejects the hypothesis of equal d across subperiods (the p-value is 0.004).

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[15] The migration figures include migration to and from abroad. In relation to the internal migration across counties migration to and from abroad is in general small. Source: Statistics Sweden, the Population Statistics, various issues. Note that per capita incomes can be seen as proxies for wage rates.

[16] I also included average yearly temperature in the county capitals (Source: SMHI (1991)) as an explanatory variable in equation (7). The estimated coefficient was positive but insignificant (the p-value was 0.23) for the whole sample period, 1911-1990. The estimates of the coefficient of real per capita income were not sensitive to the inclusion of the temperature-variable, but remained fairly constant.

To calculate b in equation (6) and subsequently b, in addition to an estimate of the sensitivity of net migration to income differentials, we also need estimates of k / k and a. Using Barro and Sala-i-Martin's (1994, ch. 9) assumptions for the U.S. States that the amount of human capital of migrants equals that of natives, that migrants do not carry physical capital (machines and buildings), that the capital share of physical capital is 0.3 and that the capital share of human capital of 0.45, implies that k / k is 0.6(=0.45/0.75). If I use the estimate of $¶m(y) / ¶ \log y$ for the whole sample period, i.e. 0.007, then the value of b is 0.0021. Given a reasonable set of values for other parameters of preferences and technology that determine the speed of convergence, this value of b implies that b is 0.0265. (For details, see appendix A). For the same parameter values but with $b = 0$ the value of b is 0.0252. Hence, for given parameter-values, migration increases the speed of convergence only marginally.

The highest estimated value of d for the subperiods is 0.027. Using this value, b increases to 0.030.

Given that $k / k < 1$, a prediction of the model is that the speed of convergence is positively related to the sensitivity of net migration to income differentials. I test this prediction by correlating the estimates of the speed of convergence for the eight subperiods with the corresponding estimates of d. I use the estimates of b from the basic regression model as well as the estimates of b from the regression model augmented with the agriculture share variable. Contrary to the prediction, both the estimated correlation coefficients are negative, although insignificant: -0.15 (-0.4) and -0.07 (-0.2) respectively. The t-statistics are given in parentheses. An outlier, the 1970s, is the cause of these negative correlations. During the 1970s trade union activity might have influenced the convergence. As a result, it can be motivated to exclude this decade. If the 1970s is excluded, the correlations are significantly positive, 0.73 (2.4) and 0.66 (2.0) respectively. Hence, this exercise provides some support for the model.

A more direct way to investigate the effect of migration on the speed of convergence is to include the average annual net migration rate, m_{i,t_0,t_0+T}, into the income growth regressions. If this is done the estimates of b should be net of the effects of migration. A problem is however that the average annual net migration rate during the sample period is not a predetermined variable. A Hausman specification test also indicates that there is a correlation between the migration variable and the disturbance term for the whole sample period when m_{i,t_0,t_0+T} is entered linearly in the basic regression model. To tackle the problem of simultaneous determination of per capita income growth and net migration, the income growth regressions are estimated by two-stage least square. Aside from the predetermined variable, initial per capita income, I use net migration rate at time t_0 and average yearly temperature as instruments for the average annual net migration rate between time t_0 and $t_0 + T$[17].

For the sample period 1919-1990 the estimate of b drops from 0.039 (9.4) to 0.032 (4.9) when net migration is held constant. This indicates that migration has a positive effect on the speed of convergence, although most of the convergence seems not to be explained by migration. For the sample period 1911-1990 the estimate of b falls by even more. This regression is however plagued by multicollinearity. The R square value of the first-stage regression would decrease only marginally, from 68.3 percent to 64.0 percent, if I were to use initial per capita income as the sole instrument, i.e almost all of the variation in the migration variable is explained by initial per capita income. Multicollinearity is less of a problem for the sample period 1919-1990[18]. As a result, one should focus on the regression results for this sample period. Moreover, the estimate of the coefficient of the migration variable is insignificantly negative. Table 4 shows the results.

[17] The assumption here is that temperature and initial net migration rate do not directly influence per capita income growth.

[18] For the sample period 1919-1990 the R square value of the first-stage regression is 66.9 percent. Table 3 shows that if the migration variable for the same period is regressed only on initial per capita income the R square value is 55.5 percent, which implies that the migration variable covaries more with the initial migration rate and temperature compared to the sample period 1911-1990. The R square values of the first-stage regression for the subperiods are: 73.7 percent for 1911-1919, 66.7 for 1919-1930, 75.4 for 1930-1940, 77.2 for 1940-1950, 82.0 for 1950-1960, 80.7 for 1960-1970, 56.6 for 1970-1980, and 53.6 1980-1990.

Those subperiods that get significant $\hat{b}$s when the basic regression model is estimated continue in general to have significant $\hat{b}$s when net migration is held constant. The exceptions are the subperiods 1930-1940 and 1940-1950. The latter subperiod is however plagued by multicollinearity. For the subperiod 1930-1940 the two-stage least square estimation does not work well. There is no clear-cut tendency for the estimates of b to fall when net migration is held constant. $\hat{b}$ increases for some subperiods and decreases for other subperiods. Moreover, the estimates of the coefficient of the migration variable is insignificant for most subperiods.

I also include average annual net migration rate in the regression model with the agriculture share variable. The results are very similar to the ones reported above. These results are displayed in table 5.

To summarize, it is found that migration has a positive, although relatively small, effect on the speed of convergence. The result that most of the b-convergence seems not be due to migration corresponds to findings for other sets of regions (see Barro and Sala-i-Martin (1994, ch. 11)). The estimate of the speed of convergence across the Swedish counties falls from almost 4 percent to around 3 percent when net migration is held constant in the income growth regressions for the whole sample period. The result that b tends to fall, when migration is controlled for, is however not robust across subperiods. Moreover, I find a positive correlation between the estimates of the speed of convergence for the subperiods and the corresponding estimated sensitivities of net migration to income differentials, as the model predicts, if an outlier, the 1970s, is excluded from the sample.

5. CONCLUDING REMARKS

This paper finds strong and robust evidence of convergence in real per capita income across the twenty-four Swedish counties 1906-1990. This evidence is not in favor of the one-sector AK-model. The evidence from the Swedish counties does however not allow us to distinguish between the two major, mutually non-exclusive, explanations of convergence: diminishing returns to capital and technological diffusion.

In accordance with other studies of convergence, regional as well as cross-country, it is found that the speed of convergence is relatively slow. The estimates of the speed of convergence across the Swedish counties are however higher than estimates obtained by Barro and Sala-i-Martin for other regional data set. Even controlling for possible institutional effects that occurred in the 1960s and 1970s on convergence such as wage compression due to trade union activity, the estimated speeds of convergence remain higher than estimates of about 2 percent per year for other sets of regions. One major likely explanation of this finding is that the current study adjusts incomes to account for regional differences in cost of living.

It can also be noted that the unconditional estimated speed of convergence for the Swedish counties is higher than conditional cross-country estimates of between 2 and 3 percent per year. Even though one should be careful in comparing unconditional and conditional estimated bs this result is consistent with an expectation that capital, labor and technology are more mobile within countries than across countries, which in general we would expect to increase the speed of convergence.

Appendix A

In this appendix migration flows are integrated into a neoclassical overlapping-generation model with labor-augmenting technological progress. The analysis draws on Barro and Sala-i-Martin (1994, ch. 9).

The total domestic population at time t , $L(t)$, is given by,

$$L(t) = L(0)e^{nt} \exp[\int_0^t m(v)dv], \qquad (1)$$

where $m(t)$ is net migration rate at time t, i.e. $m(t) \equiv M(t)/L(t)$. $M(t)$ is net migration at time t. $L(0)$, the population at time 0, is assumed to consist of families that arrived to the economy before time 0. Henceforth, $L(0)$ is normalized to one. Domestic residents at $t > 0$, i.e. descendants of natives, and immigrants and their descendants, consist of immortal families that grow at the exogenous rate n. The analysis also applies for emigration, $m(t) < 0$, given that the domestic population does not care about families that leave.

Immigrant households are indexed by their vintage $j \geq 0$ of arrival in the economy. For native families, I set $j = 0-$, i.e. they arrived before time 0. Households of each vintage j maximizes utility at time t,

$$U(j,t) = \int_t^{\infty} \{\log[c(j,v)]e^{-(r-n)(v-t)}\} dv, \qquad (2)$$

s.t. its flow budget constraint: $da(j,v)/dv = (r(v)-n)a(j,v) + w(v) - c(j,v)$ (3)
and initial assets, $a(j,j)$.

r is the subjective discount rate. $c(j,v)$ and $a(j,v)$ is per capita consumption and per capita assets for households of vintage j at time v. For immigrant households, i.e. households indexed by $j \geq 0$, I assume that $a(j,j)$ is the amount of capital they carry with them to their new home. Typically, migrants do not carry much physical capital (machines and buildings). $a(j,j)$ should therefore consist mainly of human capital. Migrants are not allowed to maintain any financial claims on foreign-source income. To simplify aggregation over immigrants of differing vintages I assume log utility. $w(v)$ is the wage rate at time v, which is equal for everyone.

I assume one type of capital, $k(t)$, that comprises both human and physical capital. Given an assumption of Cobb-Douglas technology, the economy's resource constraint (per effective worker) is given by,

$$\dot{\hat{k}}/\hat{k} = A\hat{k}^{a-1} - \hat{c}/\hat{k} - (x+n+d) - m*[1-(\hat{\kappa}/\hat{k})] \qquad (4)$$

where x is the exogenous rate of labor-augmenting technological progress, and d is the depreciation rate. $\hat{k}$ is capital per effective worker and $\hat{c}$ is consumption per effective worker. In contrast to a standard closed economy Ramsey model with no migration, the effective depreciation rate of capital is now augmented by $m*[1-(\hat{\kappa}/\hat{k})]$. $\hat{\kappa}$ is capital per effective migrant.

The optimization problem of migrants is not explicitly modeled. I just postulate that net migration is positively related to the capital-labor ratio, i.e. $m(\hat{k})$. A higher capital-labor ratio, everything else held constant, increases wages and therefore net migration. Moreover, the migration function, $m(\hat{k}) * [1-(\hat{k} / \hat{k})]$, is approximated by a log-linear function, $b * [\log(\hat{k} / \hat{k}_w)]$, i.e.,

$$m(\hat{k}) * [1-(\hat{k} / \hat{k})] = b * [\log(\hat{k} / \hat{k}_w)] \qquad (5)$$

where $\hat{k}_w$ is the average effective capital-labor ratio in other economies, which in this study can be interpreted as the average effective capital-labor ratio in the rest of the counties. If $\hat{k} = \hat{k}_w$, then no net migration takes place. The coefficient b is important for the convergence analysis. A higher b increases b. To see what b represents I differentiate (5) with respect to $\log \hat{k}$ [1],

$$b = [1-(\hat{k} / \hat{k})] * ¶m(\hat{k}) / ¶ \log \hat{k} \qquad (6)$$

Hence, b depends positively on the sensitivity of net migration to $\log \hat{k}$ and negatively on the ratio $\hat{k} / \hat{k}$. Using the production function, $\hat{y} = A\hat{k}^a$, and applying the chain rule, equation (6) can be written as,

$$b = a * [1-(\hat{k} / \hat{k})] * ¶m(\hat{y}) / ¶ \log \hat{y} \qquad (7)$$

Solving the maximization problem in equation (2), aggregating across vintages of immigrants, and using the market equilibrium condition yields an expression for the growth rate of effective per capita consumption,

$$\dot{\hat{c}} / \hat{c} = Aa\hat{k}^{a-1} - (x + r + d) - m * (r - n)(\hat{k} - \hat{k}) / \hat{c} \qquad (8)$$

Substituting equation (5) into equation (4) and (8) we have the two differential equations that describe the dynamics of the model. Except for the "migration"-terms, these differential equations are identical to that of a standard closed economy Ramsey model with log-utility.

Given that the Swedish counties share the same steady state it is plausible to assume that the steady state value of $\hat{k}$ is equal to the average effective capital-labor ratio in the rest of the counties, i.e. $\hat{k}_{ss} = \hat{k}_w$. This assumption implies that the steady state of the model is identical to that of the standard closed economy Ramsey model with log-utility. Furthermore, the model is saddle-path stable. To quantify the transitional dynamics I log-linearize the differential equations of the model around steady state. The convergence coefficient is given by,

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[1] Following Barro and Sala-i-Martin, I treat $\hat{k} / \hat{k}$ as a constant when deriving equation (6). Equation (6) is then an approximation that is satisfactory when $m(\hat{k})$ is relatively small.

$$b = 2^{-1}\left\{\left[j^{2} + 4b(r-n) + 4(1-a)(r+d+x)\left[(r+d+x)/a - (n+x+d)\right]\right]^{1/2} - j\right\} \quad (9)$$

where $j = r - n - b$.

To quantitatively assess the effect of migration on b I use equation (9) together with the parameter values, $r = 0.02, x = 0.02, d = 0.05, n = 0.005,$ and $a = 0.75$. Except the value of n, which is the average annual population growth rate for Sweden 1911-1990, these are the values used by Barro and Sala-i-Martin (1994, ch. 9).

To calculate b, I use the estimate of $\P m(\hat{y}) / \P \log \hat{y}$ for the whole sample period, 0.007. Moreover, using Barro and Sala-i-Martin's (1994, ch. 9) assumptions for the U.S. States that the amount of human capital of migrants equals that of natives, that migrants do not carry physical capital, that the capital share of human capital is 0.45, and that the capital share of physical capital is 0.3, implies that $\hat{k} / \hat{k}$ equals 0.6 (=0.45/0.75). Using these parameter-values in equation (7) implies that b is 0.0021, which gives a value of b of 0.0265. For the same parameter values but with $b = 0$, b is marginally lower, 0.025.

Appendix B: On the Adjustment of Incomes for Differences in Housing Costs

For years that lack data on housing costs I use z_t of the most adjacent year. Since the regional differences with respect to rents of adjacents years are very similar - the correlations between rents of adjacent years are all 0.93 or higher - this approach is not likely to create any problems during the first half of the sample period. The correlation between rents in 1945-1946 and sales prices of one family houses in 1952-1956 is however lower, 0.66.

Table 1 presents the results from the cross-county income growth regressions. In rows 7 and 8 of table 1 incomes in 1950 are adjusted on the basis of the average house prices for the period 1952-1956. If I instead adjust on the basis of the rents in 1945-1946 I get basically the same regression results. For the sample period 1940-1950 the estimate of b is 0.063 (7.0) when the incomes in 1950 are adjusted on the basis of the rents in 1945-1946. This estimate is to be compared with a estimate of 0.060 (7.7) reported in table 1. For the sample period 1950-1960 the estimated b is 0.022 (2.9) when the incomes in 1950 are adjusted on the basis of the rents in 1945-1946. The estimate reported in table 1 is 0.024 (3.5). Moreover, the joint estimate of b for the eight subperiods is still 0.037 (26.9) when the incomes in 1950 are adjusted on the basis of the rents in 1945-1946. The fact that one gets virtually the same regression results suggests that the use of two different non-overlapping series as proxies for housing costs is not likely to introduce any serious measurement problems.

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Table 1. Cross-County Income Growth Regressions.

Dependent variable: Average annual real per capita income growth rate.

	Adjusted Incomes			Non-Adjusted Incomes		
SAMPLE	Constant	$\hat{b}$	R^2	Constant	$\hat{b}$	R^2
1. 1906-1990	.045 (381.0)	.041 (8.1)	99.5	.044 (325.6)	.027 (15.0)	99.3
2. 1911-1990	.047 (178.7)	.037 (8.4)	99.2	.046 (190.4)	.024 (16.8)	99.2
3. 1919-1990	.052 (150.9)	.039 (9.4)	99.2	.050 (148.5)	.027 (16.2)	99.0
4. 1906-1916	.097 (25.1)	.040 (5.2)	59.2	.096 (24.6)	.039 (5.0)	58.1
5. 1919-1930	.065 (13.8)	.018 (4.2)	48.8	.061 (13.7)	.014 (3.6)	41.4
6. 1930-1940	.054 (6.5)	.021 (3.7)	39.0	.056 (6.7)	.022 (3.9)	45.7
7. 1940-1950	.154 (18.0)	.060 (7.7)	81.8	.155 (19.4)	.060 (8.3)	83.5
8. 1950-1960	.085 (6.0)	.024 (3.5)	43.6	.067 (5.5)	.015 (2.8)	35.9
9. 1960-1970	.185 (19.9)	.066 (10.8)	89.9	.110 (11.1)	.026 (6.0)	65.3
10. 1970-1980	.228 (12.5)	.093 (6.8)	81.3	.166 (14.3)	.054 (9.1)	86.3
11. 1980-1990	.025 (0.7)	.003 (0.3)	0.0	.019 (0.9)	.001 (0.1)	0.1
12. Eight periods, b restricted	--	.037 (26.9)	--	--	.032 (31.3)	--
13.Likelihood ratio statistic of equal bs (p-value)	--	32.2 (0.000)	--	--	28.3 (0.000)	--

Notes: For the whole sample period (row 1-3) the estimation method is nonlinear least square. For the subperiods the estimation technique is nonlinear SUR. The t-statistics are given in parentheses. The likelihood ratio statistic in row 13 refers to a test of equality of the convergence coefficient over the eight subperiods. The p-value comes from a c^2 distribution with seven degrees of freedom.

Table 2. Cross-County Income Growth Regressions with Agricultural Share.

Dependent variable: Average annual real per capita income growth rate.

SAMPLE	$\hat{b}$ reference	Constant	$\hat{b}$	agr_{t_0}	R^2
1. 1911-1990	.037 (8.4)	0.05 (27.6)	.033 (3.5)	.001 (0.4)	99.2
2. 1919-1990	.039 (9.4)	0.05 (27.4)	.037 (3.7)	.000 (0.2)	99.2
3. 1911-1919	.007 (1.2)	0.09 (2.8)	.034 (1.9)	-.063 (-1.7)	17.1
4. 1919-1930	.016 (3.9)	0.06 (2.8)	.017 (1.6)	.000 (0.0)	48.7
5. 1930-1940	.022 (3.9)	0.04 (1.3)	.016 (1.2)	.014 (0.5)	39.3
6. 1940-1950	.058 (7.6)	0.21 (8.8)	.108 (4.1)	-.045 (-2.6)	85.2
7. 1950-1960	.026 (3.8)	0.18 (6.7)	.076 (3.9)	-.044 (-3.4)	53.5
8. 1960-1970	.067 (10.7)	0.20 (9.6)	.074 (5.4)	-.006 (-0.7)	89.7
9. 1970-1980	.096 (6.8)	0.26 (9.2)	.124 (4.3)	-.012 (-1.1)	82.0
10. 1980-1990	-.001 (-0.1)	0.09 (1.6)	.022 (1.2)	-.039 (-1.6)	7.8
11. Eight periods, b restricted	.025 (19.1)	--	.046 (9.6)	--	--
12. Likelihood ratio statistic of equal bs (p-value)	55.4 (0.000)	--	18.0 (0.012)	--	--

Notes: The regressions with agricultural share are based on equation (4) with the counties' initial agricultural shares added linearly. Estimation method is nonlinear least squares for the whole sample period (row 1-2). For the subperiods the estimation method is nonlinear SUR. The t-statistics are given in parentheses. The likelihood ratio statistic in row 12 refers to a test of equality of the convergence coefficient over the eight subperiods. The p-value comes from a c^2 distribution with seven degrees of freedom. $\hat{b}$ reference is $\hat{b}$ from the basic regression equation, i.e. from equation (4). The subperiod in row 3 is here however 1911-1919 and not 1906-1916 as it was in table 1.

Table 3. Cross-County Migration Regressions

Dependent variable: Average annual net migration rate.

SAMPLE	Constant		Log Real Per Capita Income		R^2
1. 1911-1990	-0.01	(-6.3)	0.007	(6.3)	64.0
2. 1919-1990	-0.01	(-5.2)	0.007	(5.2)	55.5
3. 1911-1919	-0.01	(-5.5)	0.006	(4.4)	39.9
4. 1919-1930	-0.01	(-5.6)	0.008	(4.5)	35.3
5. 1930-1940	-0.02	(-5.9)	0.009	(5.7)	58.1
6. 1940-1950	-0.03	(-7.6)	0.015	(7.6)	75.2
7. 1950-1960	-0.07	(-7.0)	0.027	(6.9)	63.8
8. 1960-1970	-0.07	(-3.2)	0.025	(3.2)	33.1
9. 1970-1980	0.08	(3.5)	-0.022	(-3.4)	13.4
10. 1980-1990	0.04	(1.2)	-0.010	(-1.1)	5.7
11. Eight periods, *d* restricted	--		0.009	(9.3)	--
12. Likelihood ratio statistic of equal *d* s (p-value)	--		20.8 (0.004)		--

Notes: Estimation method is SUR. The t-statistics are given in parentheses. The likelihood ratio statistic refers to a test of equality of the regression coefficient of the logarithm of real per capita income over the eight subperiods. The p-value comes from a c^2 distribution with seven degrees of freedom.

Table 4. Income Growth Regressions with Migration

Dependent variable: Average annual real per capita income growth rate.

SAMPLE	$\hat{b}$ reference		Constant		$\hat{b}$		*migration rate*		R^2
1. 1911-1990	.037	(8.4)	0.05	(29.1)	.022	(2.8)	-0.24	(-1.2)	98.3
2. 1919-1990	.039	(9.4)	0.05	(55.6)	.032	(4.9)	-0.10	(-1.0)	99.0
3. 1911-1919	.008	(1.2)	0.04	(3.4)	.009	(0.9)	0.17	(0.2)	6.6
4. 1919-1930	.019	(4.1)	0.08	(12.5)	.027	(5.1)	0.89	(2.2)	74.4
5. 1930-1940	.021	(3.4)	0.03	(1.2)	.006	(0.4)	-1.21	(-1.1)	16.2
6. 1940-1950	.060	(7.3)	0.12	(1.8)	.034	(0.7)	-0.90	(-0.5)	74.6
7. 1950-1960	.029	(3.6)	0.11	(3.2)	.036	(2.0)	0.19	(0.5)	47.7
8. 1960-1970	.067	(9.9)	0.18	(13.6)	.065	(7.4)	-0.03	(-0.4)	90.2
9. 1970-1980	.091	(6.0)	0.24	(9.9)	.099	(5.2)	-0.17	(-1.0)	81.7
10. 1980-1990	-.001	(-0.1)	-0.01	(-0.2)	-.007	(-0.6)	0.35	(1.5)	10.6

Notes: The regressions are based on equation (4) with average annual net migration rate added linearly. Estimation method is nonlinear two-stage least squares. Aside from the predetermined variable, the logarithm of initial per capita income, I use net migration rate at time t_0 and average yearly temperature as instruments for the migration variable. The t-statistics are given in parentheses. $\hat{b}$ reference is $\hat{b}$ from the basic regression equation: the regression equations of the subperiods have been estimated individually.

Table 5. Income Growth Regressions with Migration and Agricultural Share

Dependent variable: Average annual real per capita income growth rate.

SAMPLE	$\hat{b}$ reference		Constant		$\hat{b}$		*migration rate*		agr_{t_0}		R^2
1. 1911-90	.033	(3.5)	0.04	(14.2)	.020	(2.5)	-0.16	(-1.2)	.002	(0.9)	98.8
2. 1919-90	.037	(3.7)	0.05	(21.7)	.029	(3.5)	-0.09	(-1.1)	.001	(0.3)	99.0
3. 1911-19	.039	(1.7)	0.15	(2.5)	.074	(1.5)	1.34	(1.1)	-.112	(-1.9)	20.4
4. 1919-30	.030	(2.1)	0.09	(4.6)	.035	(3.0)	0.89	(2.2)	-.015	(-0.7)	75.1
5. 1930-40	.015	(0.9)	0.00	(0.0)	-.004	(-0.2)	-1.22	(-1.0)	.024	(0.6)	16.9
6. 1940-50	.107	(3.6)	0.23	(3.1)	.135	(1.0)	0.40	(0.3)	-.046	(-2.1)	86.8
7. 1950-60	.068	(2.8)	0.17	(4.2)	.065	(2.5)	-0.15	(-0.3)	-.041	(-1.8)	52.0
8. 1960-70	.066	(4.5)	0.18	(7.6)	.066	(4.5)	-0.03	(-0.3)	-.001	(-0.1)	90.2
9. 1970-80	.111	(3.4)	0.25	(6.8)	.113	(3.4)	-0.10	(-0.7)	-.009	(-0.7)	82.4
10. 1980-90	.018	(0.8)	0.11	(1.3)	.030	(0.9)	0.25	(0.9)	-.025	(-0.7)	9.2

Notes: The regressions are based on equation (4) with average annual net migration rate and initial share of population in agriculture added linearly. Estimation method is nonlinear two-stage least squares. Aside from the predetermined variables, the logarithm of initial per capita income and initial agricultural share, I use net migration rate at time t_0 and average yearly temperature as instruments for the migration variable. The t-statistics are given in parentheses. $\hat{b}$ reference is $\hat{b}$ from the basic regression equation augmented with the agriculture share variable: the regression equations of the subperiods have been estimated individually.

Figure 1:
Average Annual Real Per Capita Income Growth Rate 1906-1990 and Log
Real (1980 Prices) Per Capita Income in 1906 (in thousands of SEK)

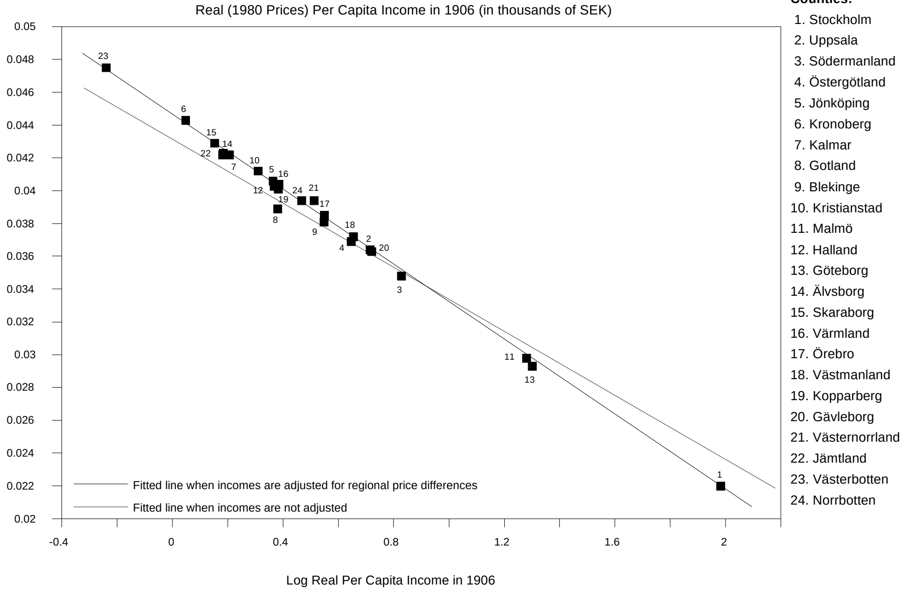

Figure 2a: Log Real (1980 prices) Per Capita Income (in thousands of SEK)

1906–90 for the Swedish Counties, Adjusted for Differences in Housing Costs

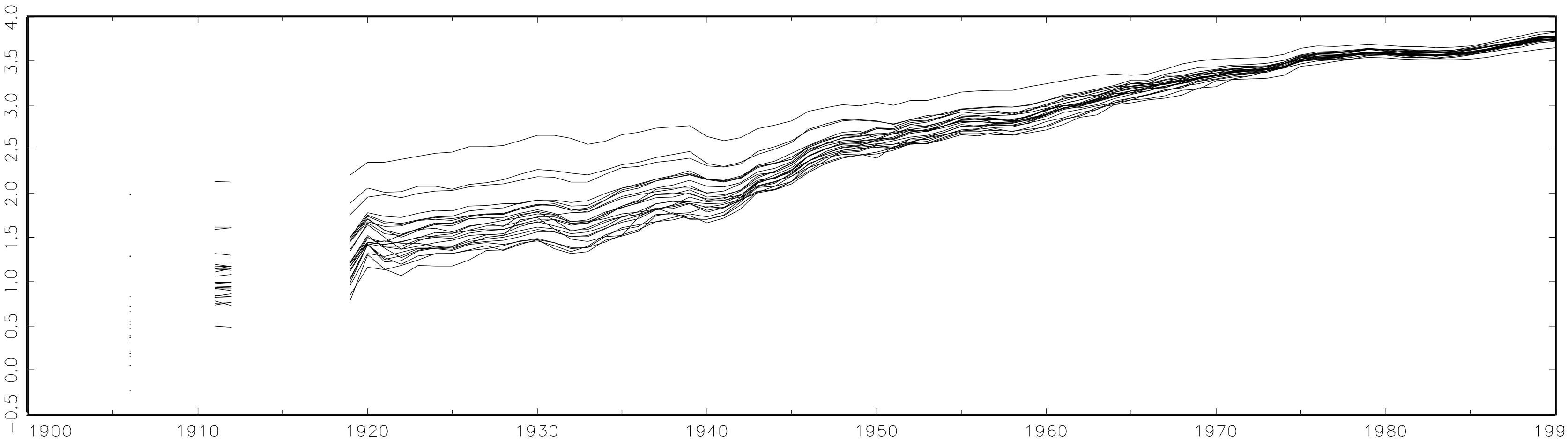

Figure 2b: Log Real Per Capita Income 1906–90 for the Swedish Counties,

Not Adjusted for Differences in Housing Costs

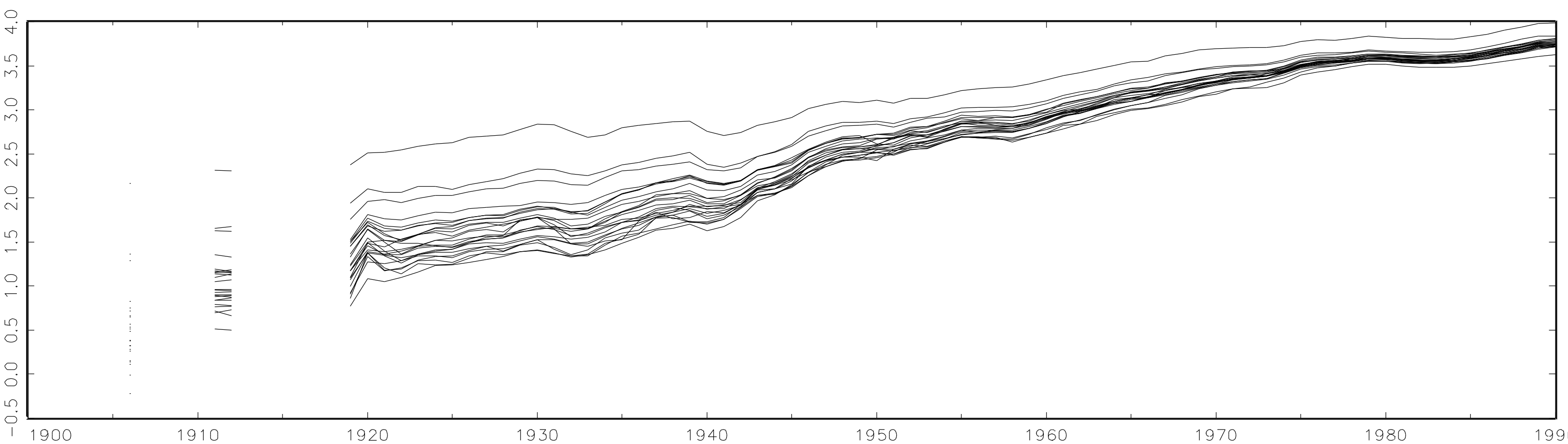

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Figure 3: Dispersion of the Logarithm of Real Per Capita Income

Across the Swedish Counties 1906–1990