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ABSTRACT

Wealth Inequality and Credit Markets: Evidence from Three Industrialized Countries*

Capital market theory predicts that the wealth distribution of an economy affects real interest rates. This paper empirically analyzes this relationship for the US, the UK and Sweden. We obtain that measures of wealth inequality are positively linked to the real rate on government securities in all three countries. This result is consistent with predictions from capital market equilibrium models with moral hazard such as Aghion and Bolton (1997) or Piketty (1997). Accordingly, rich individuals can only credibly commit to providing effort if the rate of return is not too high. When the rich are poorer, the rate of return has to be lower in order to guarantee entrepreneurial effort. Capital demand will therefore fall as inequality is reduced. The capital market is in equilibrium at a lower rate of return. The results bear important implications for economic growth and distributive policies.

JEL Classification: E20 and E62
Keywords: interest rates and wealth distribution

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1 Introduction

Understanding the role of wealth inequality on credit markets is crucial in many respects. For instance, investors need to know the impact on real interest rates when they decide on where to invest. And governments need to be aware of the mechanisms at play when they intend to implement policies targeting equity or growth concerns.

One view that has been put forward by Aghion and Bolton (1997) and Piketty (1997) is that inequality may be unproblematic when high individual wealth that goes along with inequality enables some individuals to become successful entrepreneurs. The wealth of these entrepreneurs eventually trickles down to the poor in course of the development process. This view is associated with the observation that rich entrepreneurs can credibly commit to providing effort at a higher interest rate. As a consequence of this, more inequality may be associated with a higher equilibrium real interest rate (Grüner and Schils, 2007).

Another, far more pessimistic view on wealth inequality is that it may lead to a misallocation of firms to low ability, but rich entrepreneurs. According to this view, low ability rich entrepreneurs crowd out high ability poor entrepreneurs on the capital market. With less wealth inequality, better entrepreneurs are selected. They can pay higher returns to investors. Similar results obtain in more complex setting, where both ability and effort are unobservable (see e.g. Grüner, 2003).

The purpose of the present paper is to empirically investigate which of the two views is consistent with the data from a number of industrialized countries: the US, the UK and Sweden. To capture inequality, we capitalize on rather recently published data from Wolff (2002) on the net worth held by the top 1% of wealth holders. We have chosen these three countries because of the availability of quite reliable long run time series.[1] Our econometric method follows a multivariate time

[1]For an overview of recent findings on wealth concentration also see Ohlsson et al. (2006).

series approach. As a baseline model, we consider a three-dimensional VAR process containing the variables real interest rate, wealth inequality and real per capita GDP. Our setup explicitly accounts for possible feedback mechanisms and thus addresses important issues pertaining to structural reverse causality. Dealing with these time series, however, also raises subtle issues concerning stationarity. We implement a variety of unit root tests to account for possible structural breaks and low power due to small sample size. Although these tests provide strong evidence for each variable individually containing a unit root, their linear combinations turn out to be stationary. This suggests that there may exist an equilibrium relationship between the original series on wealth distribution and real interest rates.

All in all, our results are consistent with predictions from capital market models with moral hazard and reject the prediction from the adverse selection model. This follows from estimating a three dimensional VECM model, which yields a positive and highly significant cointegration relationship between real interest rates and wealth inequality. For instance, when studying the post World War II period only, which ranges from 1948 to 1991 in our sample, we obtain the following cointegration vectors: $(1, -0.279)$for the US, $(1, -0.465)$for the UK and $(1, -0.720)$for Sweden. They are all significant at conventional confidence levels. The UK reveals the most pronounced error correcting mechanism, where a one percentage point deviation from equilibrium on average leads to a change of the real interest rate by over 0.960 percentage points in the subsequent period. Whereas for the US and Sweden, the respectively estimated loading coefficients are $-0.665$and -0.854, both highly significant at over 99 percent confidence. We also find strong evidence that a similar relationship between wealth distribution and real interest rates exist for the period from 1924 to 1948, thus including major shocks such as the Great Depression and the second World War II. Moreover, computed impulse response functions point out for all three countries that a shock, making the wealth distribution more equal, is associated with a decrease of

the real interest rate in the subsequent years. Higher interest rates in turn lead to significantly higher levels of inequality in the UK and Sweden, but show no significant effect on the wealth distribution in the USA.

Research on wealth inequality has mostly been cast in terms of its growth implications. Thereby, three theoretically motivated channels have been propagated: First, political economy considerations. According to e.g. Alesina and Rodrik (1994), Bertola (1993) or Persson and Tabellini (1994), a poorer median voter opts for higher tax rates which incur greater distortions, so hampering growth. Second, social conflicts. While e.g. Alesina and Perotti (1996) draw on reduced investment levels ensued by political instability, Rodrik (1998) blames the ability of political systems to efficiently respond to external shocks. Whereas Fajnzylber et. al. (1998) build on the high opportunity costs caused by violence. And third, capital market imperfections, upon also we draw. Especially Galor and Zeira (1993) and Banerjee and Newman (1993) established the view that with imperfect capital markets, the wealth of an agent is crucial for gaining access to credit because of its commitment value. Yet, without credit, the poor are prevented to choose the most productive activity given their skills, which affects aggregate output (through the allocation of capital) and growth.

After the refutation of the hitherto prevailing Kuznets (1955) hypothesis of inequality first increasing and then decreasing in the course of development (see e.g. Deininger and Squire, 1998) less the direction of causality from inequality to measures of economic performance, but rather the sign and significance of the relationship remain controversial issues (also see the surveys by Benabou, 1996; Ferreira, 1999; and Perotti, 1996). The study results seem to differ mainly because of the use of different data sets and econometric methodologies. Also the employed measures of income inequality turned out to be an inadequate proxy for wealth inequality. Not

to forget potential non-linearity, functional form and identification problems.[2] When empirically scrutinizing the link between inequality and growth, the three channels of influence reviewed above –partly supplementary, partly complementary in nature– interfere with each other. The paper at hand therefore focuses on the microfoundation of the capital market channel only. But as shown in greater detail below, capital market imperfections can give rise to a positive and a negative link between inequality and access to credit. Hence, it is essential to know which effect prevails. We also apply a different econometric methodology and do not rely on income inequality to proxy wealth inequality. That is why our work is most closely related to a recent empirical literature that links measures of "asset" inequality to macroeconomic variables. Besley and Burgess (2000) study the role of land reform for economic growth. They find that land redistribution reduces poverty and increases wages of the landless. Whereas Deininger and Olinto (2000) obtain a negative relationship between an unequal asset distribution and growth. The present analysis instead attempts to study how the mechanisms through which the distribution of wealth affects the capital market.

The paper is structured as follows. Section 2 first presents two simple models on the theoretically possible directions of actions. Against this background, Section 3 describes the data and the estimation method that give rise to the results in Section 4. Finally, Section 5 concludes.

2 Two theoretical arguments

In this section, we briefly explain the interest rate effects of redistribution that can be derived from capital market equilibrium models with moral hazard: first, the positive

[2]Fore more see Townsend and Ueda (2006) and Banjeree and Duflo (2005).

link in a simplified version of Aghion and Bolton (1997)[3] and then the negative link in a simplified version of Grüner (2003).

2.1 Inequality increasing the rate of return

Consider an economy populated with a continuum of potential entrepreneurs of mass 1. The wealth endowment of an agent is denoted by w_i. The cumulative distribution of wealth is denoted $\Phi(w)$, average wealth is $\bar{w}$. Each agent may start an investment project which a fixed capital outlay $I > w$. An agent who does not start a project lends money to other agents. An entrepreneur with wealth $w_i < I$ needs credit $I - w_i$from an investor. A project yields a positive return Y with probability p (resp. q) if the agent does (resp. does not) provide effort. With probability $1-p$ (resp. $1-q$) the project fails and returns zero. Effort costs B monetary units and is unobservable. All agents are risk neutral.

Consider a credit contract that yields a risk free return of R per unit lent by the investor. Under such a contract, an entrepreneur must pay $R(I-w)/p$if he succeeds, he repays nothing in case of failure. This contract induces effort if shirking does not pay, i.e. if

$$p\left[Y - R(I - w_i)/p\right] - B \geq q\left[Y - R(I - w_i)/p\right]. \tag{1}$$

Solving this incentive constraint for the agent's wealth w yields:

$$w_i \geq \omega(R) := I - \frac{p}{R}\left(Y - \frac{B}{p-q}\right). \tag{2}$$

Accordingly, entrepreneurs can commit to providing effort if they are sufficiently wealthy. Otherwise, the lender anticipates that the borrower will shirk and only

[3]See Grüner and Schils (2007), for a detailed formal analysis of the interest rate effect in such a model.

offers him a contract involving a repayment $R(I-w)/q$.

If the rate of return R satisfies

$$\Phi(\omega(R)) = 1 - \bar{w}/I, \tag{3}$$

then the wealth constraint holds exactly for a fraction $\bar{w}/I$ of individuals and the capital market clears. Suppose now that the market clears at a rate of return R which satisfies

$$\frac{qY}{I} =: \underline{R} < R \leq \bar{R} := \frac{pY - B}{I}, \tag{4}$$

so that only agents with wealth above $\omega(R)$ become entrepreneurs, while all others find it more profitable to become investors (see *Figure 1*).

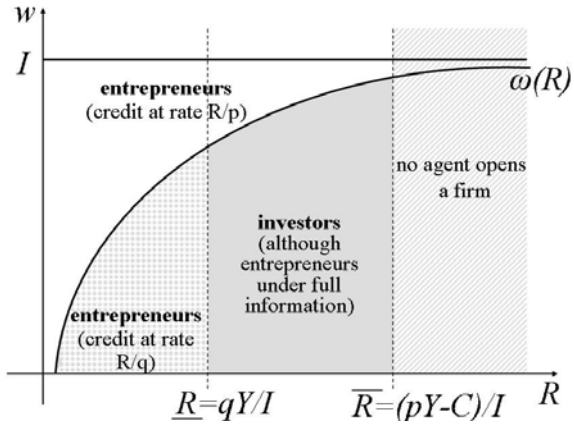


Figure 1: Solution to the individual contracting problem given R and w

Then, a more equal society is associated with poorer entrepreneurs. In equilibrium, condition (2) must hold at a lower wealth level and therefore also at a lower rate of return.[4] Entrepreneurs can only commit to providing effort if R is sufficiently low.

[4]Already in Stiglitz and Weiss (1981), the competitive equilibrium involves credit rationing, if the "Walrasian return rate" is such that there is a lower R for which the lenders' profit is higher.

This explains, why equality may be linked negatively to the rate of return R.[5]

2.2 Inequality decreasing the rate of return

The opposite effect of inequality on interests rates can be derived from a setting, where agents differ in their ability. Consider the case where ability, measured by the individual probability of success p_i, is observable and differs among agents. In this case, wealth can be used as a substitute for ability. This follows from solving the incentive constraint (1) for the interest rate R:

$$R \leq \rho(w_i, p_i) := \frac{p_i}{I - w_i}\left(Y - \frac{B}{p - q}\right). \tag{5}$$

According to this inequality, an agent can commit to providing effort only if the rate of return is not too large. As $d\rho/dp_i > 0$ and $d\rho/dw_i > 0$, a more able entrepreneur needs less wealth in order to obtain credit. An unequal wealth distribution may then be associated with low-ability rich agents who crowd out poor agents with higher ability on the capital market. The ensuing rate of return for investors may be lower than under a more equal wealth distribution, where better entrepreneurs get credit for their projects.

To see why, consider the example of an economy with two wealth classes $w_1 = \bar{w}+a$ and $w_2 = \bar{w} - a$ with $a > 0$. Half of the population are in the upper class and a measures inequality. In both classes ability levels are uniformly distributed on the interval $[p_l, p_h]$. Consider a case, where low-ability rich individuals can commit to providing effort at a rate which is slightly higher than the one for high-ability poor

[5]Interpreting more inequality as more mass above the critical collateral threshold, the depicted positive link also arises in models that aim at disentangeling the problem of information asymmetries on capital markets and that of non-convex technologies by letting the investment size vary with the market rate of return (see e.g. Piketty, 1997; Gerling, 2007).

agents, i.e.

$$\rho(\bar{w} + a, p_l) = \rho(\bar{w} - a, p_h) + \varepsilon. \tag{6}$$

Moreover, assume that there is enough capital to endow exactly 50 percent of the population with an entrepreneurial project, i.e. $\bar{w} = 1/2 \cdot I$. In such a situation, only rich agents become entrepreneurs and the equilibrium rate of return is $R = \rho(\bar{w}+a, p_l)$. Now, consider a reduction in inequality to $a' < a$ such that

$$\rho(\bar{w} + a', p_l) < \rho(\bar{w} - a', p_h). \tag{7}$$

The increase of the wealth of poor agents has enabled the high-ability poor to get credit at interest rates, at which the low-ability rich would not be able to commit to providing effort. In equilibrium, the wealth constraint is binding for high-ability poor agents and the rate of return is $R = \rho(\bar{w} - a', p_h)$. For small enough values of ε, the new equilibrium rate of return is higher than $\rho(\bar{w} + a, p_l)$, since $\rho'_w > 0$.

Hence, less inequality rises the wealth of the marginal poor entrepreneur who overtakes the rich entrepreneur with low ability in terms of the interest rate level at which he can provide effort (see *Figures 2* and 3).

Grüner (2003) has generalized this argument in a model with unobserved entrepreneurial abilities. In this model, redistribution of wealth may not only lead to the selection of better entrepreneurs and to a higher rate of return, but also to a Pareto-improvement. After redistribution, former low ability entrepreneurs receive a higher payoff from investing their reduced funds on the capital market at a higher equilibrium rate of return.

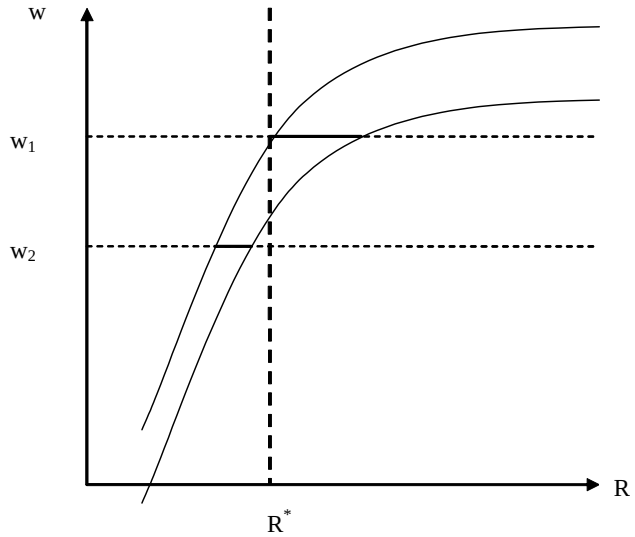


Figure 2: Before redistribution, some rich low-ability agents crowd out more qualified agents on the capital market.

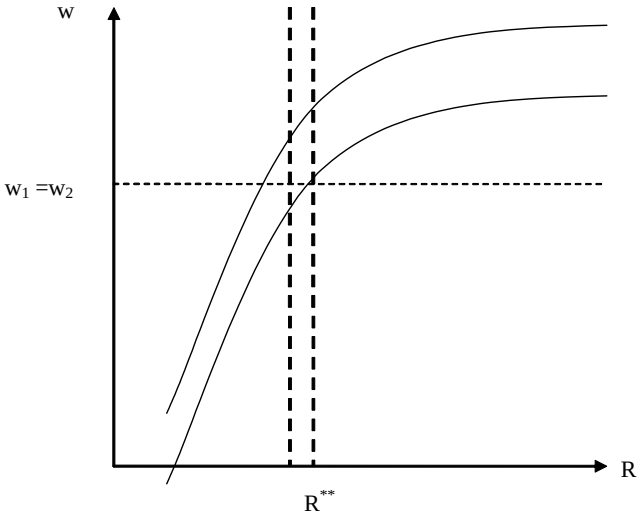


Figure 3: After redistribution, the most qualified agents become entrepreneurs and the rate of return increases from R^* to R^{**}.

3 Data

We study the relationship between wealth inequality and real interest rates for three OECD countries –the United States (1922-1992), the United Kingdom (1924-1991) and Sweden (1924-1992) – using annual data. *Table 1* provides the summary output of descriptive statistics.

3.1 Descriptive Statistics

Our measure for **inequality** is the percentage share of marketable net worth held by the top percentile of wealth holders. The data is provided by Wolff (2002) with an annual frequency for US households from 1922 to 1992, for UK adults from 1924 to 1991 and for Swedish adults from 1920 to 1992. As depicted in *Figure 4*, all three countries experienced a steady and roughly linear decline in inequality until the end of the 1970's. Inequality was particularly high in the beginning of the sample period for the UK, where over 60 percent of the net worth was held by the top percentile of wealth holders. Whereas in Sweden and the US, it was substantially lower by a difference of nearly 20 percentage points. By 1980, inequality was fairly similar across the three countries under consideration, fluctuating around a value of 20 percent for Sweden and the UK thereafter. For the US however, inequality picked up again. By 1992, over 32 percent of all marketable net wealth was held by the top percentile of wealth holders.

Real interest rates r_t are computed using the standard Fisher equation $(1 + r_t) = (1 + i_t) / (1 + \pi_t^e)$, where i_t denotes the yield on 10-year government bonds and π_t^e the expected inflation rate. We assume perfectly rational expectations and use the actual inflation rate as the expected one (i.e. $\pi_t^e = \pi_t$).[6] In order to increase robustness, our

[6] Also see e.g. Rapach and Weber (2004), Ireland (1996) or Evans (1998).

analysis relies on several measures and data source for the interest rate. In case of the US, eh.net (2006) presents data on the 3 to 6 month commercial paper rate from 1831 on, which is based on work by Macaulay (1938) and the records of the Federal Reserve Board. On top, we draw on the corporate bond rate (1870-1997), the short rate (1871-1997) and the long rate (1873-1997) published by Chada and Dimsdale (1999) as well as on the yield on 10-years government bonds (1900-2001) compiled on the basis of data published by Global Financial Data (2002) and the Federal Reserve Bank of St. Louis. For the UK, we use a time series on the 2,5% consol yield (1875-2001) derived from Global Financial Data (2002), which is based on Neal (1990) and regular publications of The Times (London), the Banker's Magazine, the Financial Times and the Central Statistical Office. Furthermore, we use yields on government bonds regularly published by the Bank of England. For Sweden, data on the yields on 10-years government bonds from 1919 to 2001 comes EcoWin (2002), Hansson and Frennberg (1992) as well as regular publications of the Swedish National Debt Office. For the inflation rates, we use the CPI data published by Officer and Williamson (2007), which is available from 1774 to 2006 for the US and from 1265 to 2006 for the UK. Whereas Statistics Sweden (2002) offers CPI data for Sweden from 1831 to 2006. Sample mean real interest rates were positive, but fluctuated fervently around the time of the Great Depression and the Second World War (*Figure 5*). During the sample period, real interest rates went as low as -10.83 percent in the US, -20.58 percent in the UK and -10.85 percent in Sweden. Sample max, on the other hand, was as high as 15.40 for the US, 23.82 percent for the UK and 8.37 percent for Sweden, with a respective standard deviation of 4.18, 6.4 and 3.63.

To explicitly control for income effects that may be correlated to both, changes in real interest rates and inequality, we include **real per capita GDP** as an additional variable. It is measured in 1990 Geary Khamis dollars and taken from Maddison (2004). As depicted in *Figure 6*, among the three countries, the US had by far the

highest average and maximum sample value of per capita GDP. Also note that after World War II, per capita GDP was similar for Sweden and the UK. However with the beginning of the 1960's, Sweden enjoyed increasingly higher per capita GDP than the UK. It is common for per capita GDP to unfold a positive and roughly linear trend.

Variable	Mean	Min	Max	Std.Dev.
Inequality_US	31.82	19.90	44.17	5.26
Inequality_UK	39.78	19.60	61.00	13.35
Inequality_Sweden	25.37	15.70	39.75	7.65
InterestRate_US	2.05	-10.83	15.40	4.18
InterestRate_UK	1.63	-20.58	23.82	6.40
InterestRate_Sweden	1.73	-10.85	8.37	3.63
GDP_US	12433.8	4777.0	23201.0	5491.6
GDP_UK	9203.6	4921.0	16430.0	3318.3
GDP_Sweden	9402.8	3130.0	17695.0	4731.6

Note: Number of observations: US (70), UK (67) and Sweden (68),

Table 1: Descriptive Statistics

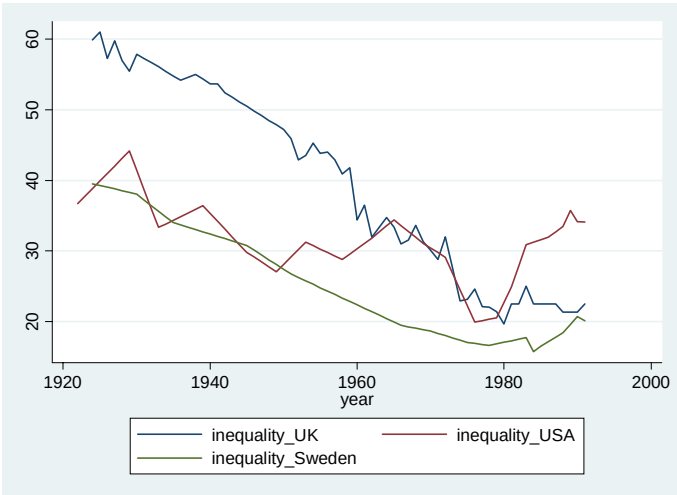


Figure 4: Share of marketable net worth held by top 1% of wealth holders (in %)

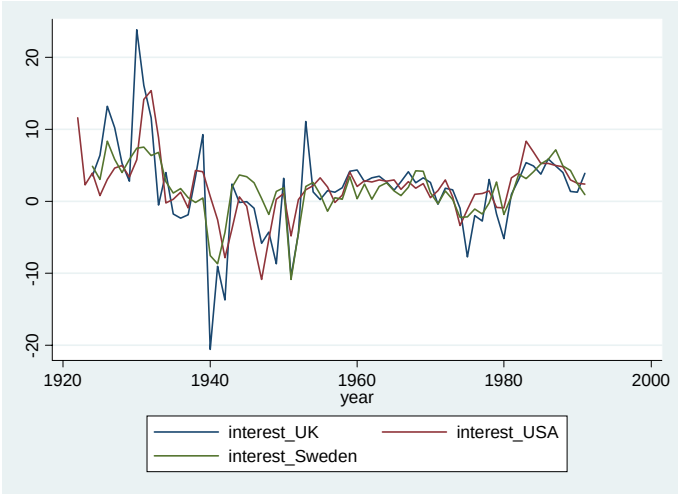


Figure 5: Real interest rates (in %)

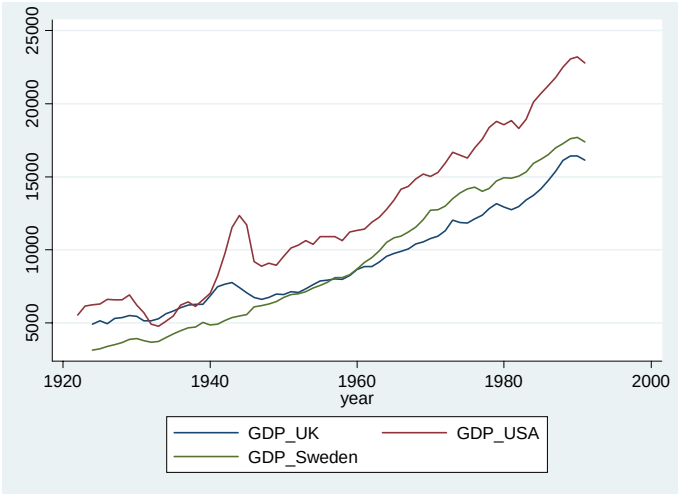


Figure 6: Per Capita GDP (in 1990 Geary Khamis Dollars)

3.2 Unit Root Tests

We apply three different unit root tests to not only address possible structural breaks, but also low power due to the small sample size. Concerning the first issue, we use

the method suggested by Lanne et al. (2002), where the data generating process may be written as:[7]

$$y_t = \mu_0 + \mu_1 t + \left[\gamma_1 (1-\theta L)^{-1} + \gamma_2 (1-\theta L)^{-1} L\right] d_t + x_t \quad t = 1, ..., T \tag{8}$$

where d_t is a shift dummy and x_t is an $AR(p)$ process with possible unit roots. For a start, the deterministic parameters μ_0, μ_1, γ_1, γ_2 and θ are estimated by GLS. Here, the timing of the structural break is chosen such as to minimize the GLS objective function. Let $\tilde{y}_t$ be the time series y_t adjusted by the obtained GLS estimates and $w_t = a(L)\tilde{y}_t$, with $a(L) = 1 - a_1 L - ... - a_p L^p$. Then, the unit root test is based on the following auxiliary regression model:

$$\Delta w_t = v + \Phi w_{t-1} + \pi a(L) \Delta d_t + \sum_{i=1}^{p} \alpha_i \Delta \tilde{y}_{t-1} + \epsilon_t \quad t = p+2, ..., T \tag{9}$$

where p is the number of lags specified for $\Delta\tilde{y}_t$ as suggested by information criteria. Under the null hypothesis of $\tilde{y}_t$ containing a unit root, Φ should be zero. The corresponding test statistics are denoted LLS and critical values can be found in Lanne et al. (2002).

For the case of a structural break occurring in the mean only, we use the Clemente et al. (1998) test procedure for additive outliers. In a first step, the deterministic part is removed by estimating:

$$y_t = \mu_0 + a_1 d_{1,t} + a_2 d_{2,t} + x_t \quad t = 1, ..., T \tag{10}$$

In a second step, we let $\tilde{y}_t$ be the time series y_t that is adjusted by the deterministic parts of equation (10). Then, we estimate the subsequent model:

$$\tilde{y}_t = \sum_{i=0}^{k} \omega_{1,i} \Delta d_{1,t-1} + \sum_{i=0}^{k} \omega_{2,i} \Delta d_{2,t-1} + \Phi \tilde{y}_{t-1} + \sum_{i=0}^{k} c_i \Delta \tilde{y}_{t-1} + e_t \tag{11}$$

[7] We consider here the case of a rational shift function. Results are robust to imposing other functional relations such as e.g. an exponential shift function.

Under the unit root hypothesis, Φ should be equal to one. The timing of the break points is in turn determined by searching for the minimal t-value for Φ. The corresponding test statistic is called CMR. Critical values can be found in Clemente et al. (1998).

As a robustness check that accounts for situations, in which the above tests may suffer from low power, we use a test proposed by Kwiatkowski et al. (1992). It is constructed under the null hypothesis that the time series is stationary. The corresponding test statistic is:

$$KPSS = \frac{\sum_{t=1}^{T} S_t^2}{T^2\sigma^2} \text{ with } S_t^2 = \sum_{j=1}^{t} x_t \text{ and } \sigma^2 = T^{-1}\sum_{t=1}^{T} x_t^2 + 2T^{-1}\sum_{s=1}^{l} w(s,l) \sum_{t=s+1}^{T} x_t x_{t-s}$$

where σ is a spectral estimator for the long-run variance of x_t, using the Barlett window $w(s,l) = 1 - s/(1+l)$. The truncation lag in turn is determined by the rule $l = int[j(4T/100)]^{0.25}$, with $j = 4, 12$.

Table 2 provides the summary output of the LSS and $KPSS$ unit root tests. *Figure 4* and *Figure 6* clearly show that inequality and GDP contain a (linear) trend. Thus, we model GDP and inequality as trend stationary. For all three countries, LSS, CMR and $KPSS$ tests unanimously indicate that these variables contain a unit root, at least during the sample period of interest. Taking first differences, the LSS and CMR reject that Φ is equal to zero at conventional confidence levels. Contrariwise, the $KPSS$ cannot reject the null hypothesis that the transformed variables are stationary.

Concerning the real interest rate, the unit root test results are not that clear. While for the US, LSS and CMR tests cannot reject the null hypothesis of a unit root in the real interest rate, the $KPSS$ rejects stationarity at the 95 (resp. 90) percent confidence level for the case, where a truncation lag of 4 (resp. 12) was used.

But significance drops to the 85 percent confidence level, when alternatively, a truncation lag of 11 was used. A similar picture arises for the UK. For Sweden, the LSS and CMR tests provide even some weak evidence for the real interest rates

being stationary. We interpret these findings as consistent with the contemporary literature, where stationarity of the real interest rate remains a controversial issue (see e.g. Rapach and Weber, 2004). Note that when we compute unit root tests for the case of first differences, we find very strong evidence that the transformed series is stationary: both, the LSS and CMR test reject the null hypothesis of non-stationarity at over 99 percent confidence, whereas $KPSS$ tests are insignificant.

	US	**UK**	**Sweden**
GDP	LLS_3: -2.207 CMR_2: -4.138 $KPSS_4$: 0.2394*** $KPSS_{12}$: 0.1688**	LLS_2: -2.116 CMR_2: -2.995 $KPSS_4$: 0.1758** $KPSS_{12}$: 0.1532**	LLS_2: -0.878 CMR_2: -2.689 $KPSS_4$: 0.3148*** $KPSS_{12}$: 0.1659**
ΔGDP	LSS_1: -5.488*** CMR_2: -7.381*** $KPSS_4$: 0.1617 $KPSS_{12}$: 0.2501	LLS_1: -4.524*** CMR_1: -4.993*** $KPSS_4$: 0.1634 $KPSS_{12}$: 0.2792	LLS_1: -4.919*** CMR_2: -6.146*** $KPSS_4$: 0.3004 $KPSS_{12}$: 0.2469
Inequality	LLS_5: -2.540 CMR_2: -3.762 $KPSS_4$: 0.1489** $KPSS_{12}$: 0.1129*	LSS_0: -1.843 CMR_2: -5.417* $KPSS_4$: 0.2202*** $KPSS_{12}$: 0.1426**	LLS_4: -2.702 CMR_2: -2.996 $KPSS_4$: 0.3124*** $KPSS_{12}$: 0.1734**
ΔInequality	LLS_0: -3.700*** CMR_1: -4.634*** $KPSS_4$: 0.1094 $KPSS_{12}$: 0.1296	LSS_0: -8.570*** CMR_1: -8.586*** $KPSS_4$: 0.2573 $KPSS_{12}$: 0.3302	LLS_5: -2.591* CMR_1: - 4.124*** $KPSS_4$: 0.6907 $KPSS_{12}$: 0.4566
Real Interest Rate	LLS_4: -1.604 CMR_2: -3.490 $KPSS_4$: 0.1763** $KPSS_{12}$: 0.1167*	LLS_2: -2.455 CMR_2: -3.169 $KPSS_4$: 0.1122* $KPSS_{12}$: 0.0916	LLS_0: -3.061** CMR_2: -5.507* $KPSS_4$: 0.177** $KPSS_{12}$: 0.1355*
ΔReal Interest Rate	LSS_2:-8.085*** CMR_1: -8.520*** $KPSS_4$: 0.1435 $KPSS_{12}$: 0.2270	LLS_1: -6.890*** CMR_1: -10.314*** $KPSS_4$: 0.0671 $KPSS_{12}$: 0.1282	LSS_2: -3.132** CMR_1:-6.314*** $KPSS_4$: 0.1143 $KPSS_{12}$: 0.1986

Note: Critical Values for the KPSS test assuming level stationarity are 0.347 (10%), 0.463 (5%), 0.739 (1%) and 0.119 (10%), 0.146 (5%), 0.216 (1%) when assuming trend stationarity. For the KPSS, the subscript number denotes the trunction lag. Critical values for the LSS test assuming trend stationarity are -2.76(10%), -3.03 (5%), -3.55 (1%) and -2.58 (10%), -2.88(5%), -3.48(1%) when assuming level stationarity. For the LSS, the subscript number denotes the specified number of autoregressive lags, as suggested by information criteria. Critical values for the CMR test with two structural breaks are -5.37 (10%), -5.70 (5%), -6.50 (1%). For the CMR, the subscript number denotes the specified number of break points. GDP, inequality and the real interest rate are modelled under the null as trend stationary, while ΔGDP, Δinequality and the Δreal interest rate were modelled as level stationary. Number of observations: US (69), UK (44) and Sweden (67).

Table 2: Unit root test

3.3 Cointegration Analysis

In the next step, we assess whether GDP, inequality and the real interest rates are cointegrated. We apply system cointegration tests, because we do not want to assume weak exogeneity of particular regressors. The implemented test procedure is based on Johansen et al. (2000) and allows for up to two structural breaks. Assuming that $X = (Interest\ Rate,\ GDP,\ Inequality)^T$ follows a $VAR(p)$ process, the $VECM$ representation may then be written as:

$$\Delta X_t = \upsilon + \alpha \left[\beta^T X_{t-1} - \tau (t-1) - \theta d_{t-1}\right] + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-1} + \Delta d_t + u_t \quad t = p+1, ..., T \tag{12}$$

where υ is a $[3x1]$ intercept, α a $[3xr]$ vector containing loading coefficients and β^T a $[rx3]$ vector containing the cointegration relationship. θ and τ are $[3x1]$ parameter vectors, d_t shift dummies modelling a break in the long-run relationship and Δd_t impulse dummies. The error term is u_t ˜$(0, \Sigma)$. GDP and our measure for inequality exhibit a linear trend. That is why a trend term is included in (12). Yet, we restrict it to the cointegration relationship in order to rule out quadratic terms. In order to determine the number of cointegration relations, the Johansen et al. (2000) procedure estimates the model in equation (12) by maximum likelihood and applies a reduced rank procedure by testing the hypothesis $H_0(r_0)$: $rk(\alpha\beta^T) = r_0$ versus $H_1(r_0)$: $rk(\alpha\beta^T) > r_0$. The likelihood ratio test statistic is computed using the estimated eigenvalues of $\alpha\beta^T$.

Table 3 provides a summary of the obtained p-values when performing the Johansen trace test, where we specified structural breaks to occur during the time of the Second World War and during the period of the Great Depression. The tests provide strong evidence that between GDP, real interest rates and inequality, there exist two cointegration relationships. The full sample estimates are listed in columns 1 to 3, whereas pre- (resp. post-) World War II estimates are listed in columns 4 to

6 (resp. 7 to 9). For the US and Sweden, the Johansen trace tests rejects the null hypothesis of only one cointegration relationship with over 99 percent confidence, but accepts the null hypothesis of two cointegration relationships at conventional confidence levels. Similarly, the trace test detects two cointegration relations for the UK time series during the periods 1924-1948 and 1948-1991.

	Full Sample			1924-1948			1948-1991		
	US (1)	**UK** (2)	**Sweden** (3)	**US** (4)	**UK** (5)	**Sweden** (6)	**US** (7)	**UK** (8)	**Sweden** (9)
0	0.0006***	0.0506**	0.0000***	0.0000***	0.0001***	0.0000***	0.0000***	0.0000***	0.0028***
1	0.0098***	0.5322	0.0062***	0.0027***	0.0625*	0.0038***	0.0009***	0.0376**	0.0469**
2	0.3396	0.6438	0.3126	0.3725	0.4023	0.2658	0.2096	0.6792	0.1323
No. of observ.	69	66	67	22	23	23	35	36	41

Note: * Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.
Listed numbers are p-values. The lag length was determined using information criteria. All estimates were computed including a linear trend and a constant in the cointegration relationship.

Table 3: Cointegration Tests

Dependent Variable: First Difference Real Interest Rate

	Full Sample			1924-1948			1948-1991		
	US (1)	**UK** (2)	**Sweden** (3)	**US** (4)	**UK** (5)	**Sweden** (6)	**US** (7)	**UK** (8)	**Sweden** (9)
Loading Coefficient (α)	-0.554*** (-6.044)	-0.88*** (-7.360)	-0.574*** (-4.293)	-0.780*** (-4.158)	-1.344*** (-7.365)	-0.753*** (-3.911)	-0.665*** (-3.800)	-0.960*** (-5.170)	-0.854*** (-4.237)
Interest Rate	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
Inequality	-0.513*** (-4.029)	-0.502 (-1.276)	-0.765*** (-2.815)	-3.144*** (-5.242)	-4.137*** (-2.696)	-4.571*** (-3.721)	-0.279*** (-2.999)	-0.465* (-1.841)	-0.720** (-2.502)
No. of observ.	69	66	67	25	23	23	42	42	42

Note: *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.
t-values are provided in brackets. The lag length was determined using information criteria as well as a set of misspecification tests. All estimates were computed including a linear trend and a constant in the cointegration relationship, with the exception of the UK during the period 1924-1991, where the constant was insignificant in the cointegration relationship.

Table 4: VECM Cointegration Estimates: Inequality and Real Interest Rates

4 Main Results

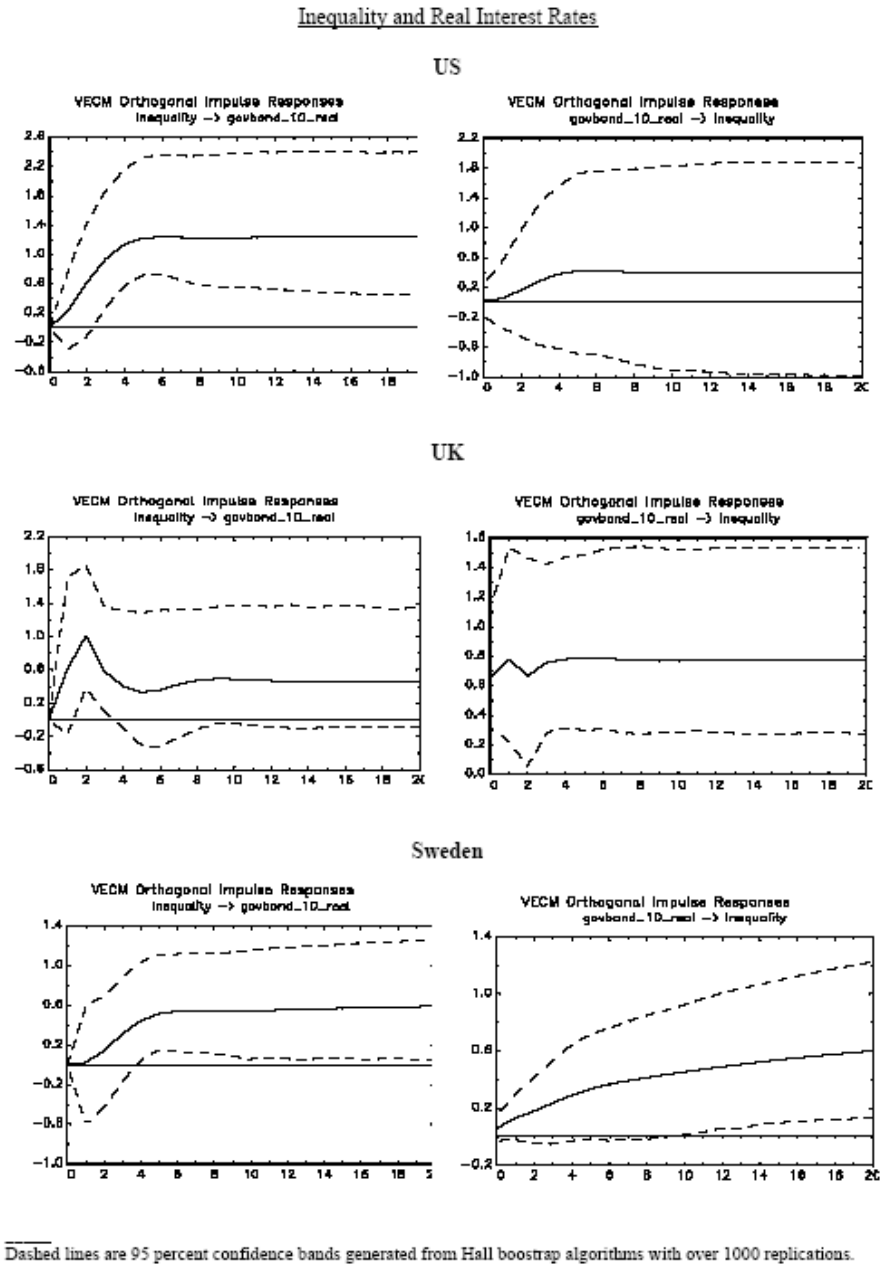
Table 4 presents the summary output of the cointegration estimates from a three-dimensional $VECM$ model as specified in equation (12), where we imposed a cointegration rank of two, as suggested by the preceding cointegration tests. With the estimate of α and β^T not being unique, we follow standard procedures and set the coefficient for the real interest rate equal to one. In turn, the lag structure was determined by using information criteria and a variety of misspecification tests. For all three countries, the 'best' model turned out to be the corresponding $VAR(2)$ process. Both, a shift dummy for World War II and an unrestricted impulse dummy for the Great Depression were included in the cointegration relationship. The shift in the cointegration relation proved to be significant for all three countries. However, this was not always the case for the impulse dummies reflecting additional temporary shocks (such as e.g. periods during which the country was involved in a major war or times of severe adverse economic conditions, such as the Great Depression).

Column 1 in *Table 4* lists the estimated cointegration relationship between the real interest rate and inequality for the US during the period 1922-1992. Both, the loading coefficient and the cointegration vector are highly significant at over 99 percent confidence. The long-run relationship is estimated as $(1, -0.513)$ with a loading coefficient of -0.554. Similar effects are found for Sweden during the period 1924-1992: the cointegration vector is $(1, -0.765)$ with a loading coefficient of -0.574, which are both individually significant at over 99 percent confidence (column 3). The estimated cointegration for the UK exhibits the largest error correcting mechanism with an estimated loading coefficient of -0.880 (column 2). However, note that even though similar in size to the estimates obtained for the US and Sweden, the cointegration vector of $(1, -0.502)$ reveals a t-statistic of -1.276, which is not significant at conventional confidence levels. Nevertheless, once the sample is separated into pre-

and post-World War II observations (see columns 5 and 7), the cointegration vector obtained for the UK is $(1, -4.137)$ and $(1, -0.465)$, which is significant at the 99 and 92 percent confidence level. When comparing estimates of the loading coefficients in columns 4 to 9, a deviation from the equilibrium relation yields the largest correcting mechanism in the real interest rate for the UK, in particular for the period 1924-1948.

Figure 7 depicts the impact that a shock in inequality exhibits on the real interest rate from othorgonalized impulse response functions, where 95 percent confidence bounds were generated using a bootstrap algorithm.[8] For all three countries, a positive shock in inequality increased the real interest rate with the impact of the marginal effect becoming steady after about 4 periods. With our multivariate time series framework explicitly allowing to address bi-directionality, we computed orthogonalized impulse response functions for the impact that shocks in the real interest rate exhibit on inequality. Interestingly, there is strong evidence that an increase in real interest rates significantly increases the share of net worth held by the top percentile of wealth holders for the UK and Sweden. Although the effect in the US is positive again, it is statistically insignificant. Also note that the generated impulse responses indicate that interest rate shocks exhibit permanent, rather than just temporary effects on the relative composition of wealth held within these economies.

[8]Dashed lines are 95 percent confidence bands generated from Hall boostrap algorithms with over 1000 replications.



Dashed lines are 95 percent confidence bands generated from Hall boostrap algorithms with over 1000 replications.

Figure 6: Impulse reaction functions

In *Table 5*, we provide a summary of the estimated cointegration relationship between inequality and real interest rates, when using short-term interest rates on deposits instead of yields on 10-year government bonds (column 1, 3 and 5).[9] Similar to the above results, we find a positive and significant cointegration relationship between the share of net worth held by the top percentile of wealth holders and the real interest rate. For the US, the cointegration relationship when using short-term (resp. long-term) deposit rates is $(1, -0.295)$ (resp. $(1, -0.258)$), and for the UK it is $(1, -0.358)$ (resp. $(1, -0.806)$).[10] Note that for both countries, the adjustment mechanism, when deviating from the long-run equilibrium, is fastest for the case of short-term interest rates.

As an additional robustness check, we computed cointegration estimates for the relation between inequality and real interest rates, when the latter are based on inflationary expectations other than perfectly rational ones (i.e. $\pi_t^e = \pi_t$, as done before). Following common practices, r_t was again generated according to the formula $(1 + r_t) = (1 + i_t) / (1 + \pi_t^e)$ with $\pi_t^e = 0.5\pi_t + 0.3\pi_{t-1} + 0.2\pi_{t-2}$.[11] Column 2, 4 and 6 of *Table 5* contain the respective summary output. For all regressions, inequality and expected real interest rates reveal a positive long-run relation, which is significant at conventional confidence levels for most of the specifications. Nevertheless, we find that the adjustment mechanism is slower for the case of expected real interest rates than the benchmark case of real interest rates.

[9] Source: Maddison (2004).

[10] Unfortunately, we were not able to obtain data on short and long-term deposit rates for the entire period 1948-1992 in the case of Sweden.

[11] When experimenting with different methods, Chadha and Dimsdale (1999) found little difference between employing complicated inflation prediction models and a simpler use of realized inflation series over the 20th century for a set of countries including the US and the UK. Backward-looking moving averages can also be found in e.g. Blanchard and Summers (1984) or Jenkinson (1996).

	Govbond	Govbond_ exp	Interest_ Long-Term	Interest_ Long-Term_exp	Interest_ Short-Term	Interest_ Short-Term_exp	Corporate Bond	Commer-cial Paper Rate
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				US (1948-1991)				
Loading	-0.665***	-0.585***	-0.652***	-0.408***	-1.009***	-0.718	-0.642***	-1.069***
Coefficient	(-3.800)	(-4.092)	(-4.473)	(-3.753)	(-5.537)	(-5.292)***	(-4.433)	(-5.877)
Interest Rate	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
Inequality	-0.279***	-0.119	-0.295***	-0.197**	-0.258***	-0.078	-0.294***	-0.230***
	(-2.999)	(-1.261)	(-3.528)	(-2.268)	(-3.217)	(-0.807)	(3.483)	(-2.797)
				UK (1948-1991)				
Loading	-0.960***	-0.367***	-0.988***	-0.382***	-0.593***	-0.287***		
Coefficient	(-5.170)	(-4.700)	(-5.367)	(-4.926)	(-3.814)	(-4.099)		
Interest Rate	1.000	1.000	1.000	1.000	1.000	1.000		
Inequality	-0.465*	-0.540*	-0.358	-0.369	-0.806**	-0.637		
	(-1.841)	(-1.649)	(-1.258)	(-1.078)	(-2.147)	(-1.527)		
				Sweden (1948 -1991)				
Loading	-0.854***	-0.459***						
Coefficient	(-4.237)	(-3.487)						
Interest Rate	1.000	1.000						
Inequality	-0.720**	-0.639**						
	(-2.502)	(-2.098)						

Note: * Significantly different from zero at 90 percent confidence, ** 95 percent confidence and *** 99 percent confidence.
The dependent variable is the first difference of the corresponding real interest rate, as indicated by the column number. T-values are provided in brackets. The lag length was determined using information criteria as well as a set of misspecification tests. All estimates were computed including a linear trend and a constant in the cointegration relationship.

Table 5: Robustness: Govbond vs. Other Real Interest Rates

5 Discussion

While theory predicts an effect of the wealth distribution on interest rates via aggregate credit supply and demand, the sign of the effect depends on the type of the underlying information asymmetry. The evidence in this paper is clearly in favor of theories based on moral hazard that link inequality to higher rates of return. They so refute theories based on adverse selection that predict a negative relationship.

Consequently, with higher real interest rate reflecting a more efficient allocation and use of capital, aggregate output and so economic growth should be higher in countries characterized by higher inequality. Barro (1999) obtains a significant negative effect of inequality on growth for poor countries, which vanishes above a certain per capita income level. This was taken as lending support to the fact that capital market imperfections are worse in poor countries, where capital markets are less developed. Against the background of this paper, the result might also be due to the fact that poor countries tend to have less aggregate wealth and wealth distributions with less mass above the crucial threshold of collateral. This also shows that one must be precautious with the meaning of inequality (we interpret it as more mass above the critical wealth level). Distributive policies reducing the wealth held by the rich might therefore turn out counterproductive with respect to long run inequality and efficiency. As less rich entrepreneurs reduce effort and thus expected project output, not only aggregate output and in turn capital accumulation over time might decrease, but also the real interest rate paid to investors. This prevents the poor from growing richer over time. They need longer to pass over the critical wealth level, which would make them eligible for credit and so empowered to become entrepreneurs themselves.[12]

[12]Banjeree and Duflo (2005) even see the endogeneity of the interest rate at the core of poverty traps, since the poor inflict a pecuniary externality on each other: low interest rates hinder the poor

Our findings also gives some support to politico-economic theories that explain limits to redistribution via an interest-rate effect on the credit market (see Grüner and Schils, 2007). This interest rate effect of inequality may explain why some middle class individuals oppose political redistribution of wealth despite the fact that they may gain wealth. Middle class individuals have incentives to oppose redistribution when the interest rate effect of redistribution dominates the wealth effect.[13]

There are also alternative explanations for the link between inequality and rates of return. In particular, one may argue that exporting capital involves some fixed costs. When only rich individuals can afford paying this cost, more inequality may be associated with more capital being exported and therefore with a higher rate of return. The analysis of individual or household data should help to shed light on this issue.

Further studies along the lines of this paper might be useful in predicting returns on investments. Studies such as those of Barro (1999) inspire to scrutinize whether our result might be confined to developed countries. These usually suffer of a higher degree and eventually different nature of capital market imperfections and display less mass above the critical wealth level. Related to this, with capital markets being increasingly integrated, the national wealth distribution may become less important while the international distribution should dominate (see Gerling, 2007). Controlling for the development of the financial system and for banking market power may be necessary in such an extended analysis.

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to save up to no longer be credit-rationed and thus to escape poverty.

[13]See Benabou (2000), Corneo and Gruener, (2000, 2002), Roemer (1998) and Piketty (1995), for other recent theoretical explanations for limits to redistribution in presence of inequalities.

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