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The Kuznets curve of the rich

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ABSTRACT

A long-standing interest in the relationship between inequality and sustainable growth continues to fascinate economists among other social scientists. It must be noted, however, that most empirical efforts have focussed on the income inequality–growth nexus, while studies on wealth inequality are much scarcer. This study attempts to fill such a gap in the literature by assessing the correspondence between the top 1 percent's wealth share and economic growth. Employing time series cointegration techniques, we study the experience of France and the United States from 1950 to 2014. Our estimates suggest that the output growth rate is an inverted-U-shaped function of the wealth share of the top 1 percent. The estimated relationship is robust to variations in control variables and estimation methods. We compute the local *optimal wealth share*, understood as the share of wealth compatible with the maximum growth rate, and show that France is growing close to its long-run potential, while the United States is significantly below its.

1. Introduction

Since ancient times, societies have been concerned about the effects of inequality on peace and prosperity. Socrates, for example, in his dialogue with Adeimantus, as reproduced in *The Republic*, demonstrates awareness of the pervasive consequences of indiscriminate wealth in deteriorating peace and order. Aristotle, in *Rhetoric*, also presents inequity as a source of conflict and anger. From Adam Smith to Karl Marx, there has been a long-standing interest in the relationship between income–wealth distribution and sustainable growth, and this continues to fascinate economists among other social scientists.

It must be noted, however, that most empirical efforts have focussed on the correspondence between *income* inequality and growth, while studies on *wealth* inequality are relatively scarce. This is rather surprising given the evidence showing that global wealth-holding is far more concentrated than income (Davies et al., 2011), and that most theoretical contributions highlight the importance of both dimensions. To the best of our knowledge, Bagchi and Svejnar (2015) and Islam and McGillivray (2020) are the only two existing references to directly assess this issue. The first relies on a *proxy* for wealth distribution from Forbes list of billionaires. The second utilises recent data published by Credit Suisse for a sample of 45 countries between 2000 and 2012. This study aims at advancing this issue by investigating the long-term correspondence between the wealth share allocated to the top 1% and output growth in France and the United States (US).

For three decades, the debate vis-à-vis rising inequality in the United States and Europe has centred on the wage premium for certain types of labour. In recent years, however, there has been an increasing realisation that most of the action is at the very top of the

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distribution (e.g., Alvaredo et al., 2013, 2018). The development of the *World Inequality Database* has facilitated the analysis and international comparison of the evolution of top income and wealth shares. The top 1 percent wealth share–growth relationship is not just a *proxy* for the inequality–growth nexus; rather it is an interesting tax policy issue in its own right. Piketty and Zucman (2015) and Alvaredo et al. (2017) have shown that the share of inherited wealth in several developed countries is back to 50–60% of the total wealth, with potentially important political and economic implications.

In 2018, the average net personal wealth of the so-called top 1 percent was 5.8 million US dollars (US$) in France and approximately US$ 12 million in the US, while its average pre-tax income was US$ 500 thousand in the former country and US$ 1.4 million in the latter. While it is certainly not a homogeneous group, such elites arguably wield a significant amount of power and have a particular set of skills that may justify their social position. This means that a very small or else a very large wealth share might be detrimental to economic performance. On the one hand, the former aforementioned case could indicate insufficient skill premia, leading to an inefficient allocation of resources. On the other hand, high wealth concentration might weaken the social consensus required to sustain growth, thereby inducing the consolidation of equally inefficient institutional arrangements in both public and private sectors.

From the interaction of these two forces, the Kuznets curve of the Rich arises as an inverted-U relationship between the output growth rate and the wealth share of those at the very top of the distribution. The original Kuznets (1955) curve was widely used to describe the association of growth and inequality over the second half of the 20th century, though it fell out of favour in recent decades (see Milanovic, 2016). This paper provides an empirical contribution to the literature by applying time series cointegration techniques to study the experiences of France and the US between 1950 and 2014. The choice of working with these two economies is twofold motivated: first, they virtually represent two different institutional arrangements in developed countries, i.e., the Anglo-Saxon and continental Europe. While recognising that there is still significant heterogeneity in the Global North (e.g. Ranaldi and Milanovic, 2021), our understanding is that their differentiation may in any case provide some important insights. Furthermore, data are available for a sufficiently long period, thereby enabling the performance of time series analysis.

Our estimates suggest that the output growth rate follows an inverted-U-shaped function of the wealth share of the so-called top 1 percent. The obtained relationship is robust to variations in control variables and estimation methods, having some similarities with Banerjee and Duflo's (2003) results. The inverted-U is also robust to narrowing the analysis to the top 0.01 percent, though it is no longer statistically significant for France. It is possible for us to compute a local *optimal wealth share*, which is defined as the share of wealth controlled by the top 1 percent and is compatible with the maximum growth rate. Any divergence from this level, in whatever direction, is associated with a reduction in long-run growth. We show that France is growing closer to its long-run potential, while the US is 1–2% below its.

As indicated by Piketty (2020), once we accept that private property will continue to play a role in society, it becomes essential to develop institutional arrangements capable of preventing the unlimited concentration of ownership which does not serve the general interest. The extreme concentration of wealth observed in European societies up to the early twentieth century did not serve to its general interest. Conversely, signs of excessive wealth concentration exacerbated social and nationalist tensions, thereby blocking educational and social investments that facilitated the development of the post-war model. At least, regarding the US, we show that the over-concentration of wealth in top percentiles of the population is damaging growth.

The remaining of the paper is organised as follows: in the next section, we briefly provide an overview of the literature on growth and distribution. Section 3 describes the main trends regarding the evolution of top wealth shares in France and the US, showing how wealth inequality is arguably more important than income inequality. Section 4 presents our empirical exercise towards identifying the inverted-U relationship. In section 5, we explore the robustness of our findings to alternative specifications and estimators. Some final considerations follow.

2. An overview of the related literature

Different mechanisms that link distribution and growth have been identified. The classical literature, for instance, has emphasised that increasing inequality might actually foster growth by channelling resources towards individuals with higher savings propensities, which consequently intensifies capital accumulation (Kaldor, 1957; Goodwin, 1967). Other scholars have suggested that concentrating income and wealth may promote the realisation of higher return projects and thus stimulate R&D investments (see, for example, Foellmi and Zweimuller, 2006; Aghion et al., 2019; Law et al., 2020). On the other hand, a large body of scholarship has indicated that increasing inequality hampers growth because it promotes institutions that prevent the protection of property rights, weakens social consensus required to sustain growth, and induces the consolidation of an inefficient state bureaucracy (Persson and Tabellini, 1994; Rodrik, 1999; Acemoglu et al., 2011). Moreover, under imperfect credit market conditions, Galor and Zeira (1993) argue that income and wealth distribution determine the allocation of human capital across individuals. Unequal societies might end up excluding certain individuals from human capital augmenting investments, which foments adverse implications in terms of growth rates.

Such a correspondence, however, does not have to be necessarily monotonic. Bhaduri and Marglin (1990) suggest that depending on how capital accumulation responds to profitability, changes in income distribution may have a positive or negative impact on growth. From a different perspective, Galor and Moav (2004) report that inequality favours growth in the early stages of development, as the economy needs to accumulate physical capital, but it is detrimental later on, when human capital becomes the critical variable.

Empirically, data availability has been a long-standing major concern. For approximately two decades, the literature was dominated by the dataset collected by Deininger and Squire (1996). More recent studies have either relied on the *World Income Inequality Database* provided by the World Institute for Development Economics Research or the *Standardized World Income Inequality Database* developed by Solt (2009), Solt (2020). Evidence generally seems to support the view that income inequality impedes growth (for a

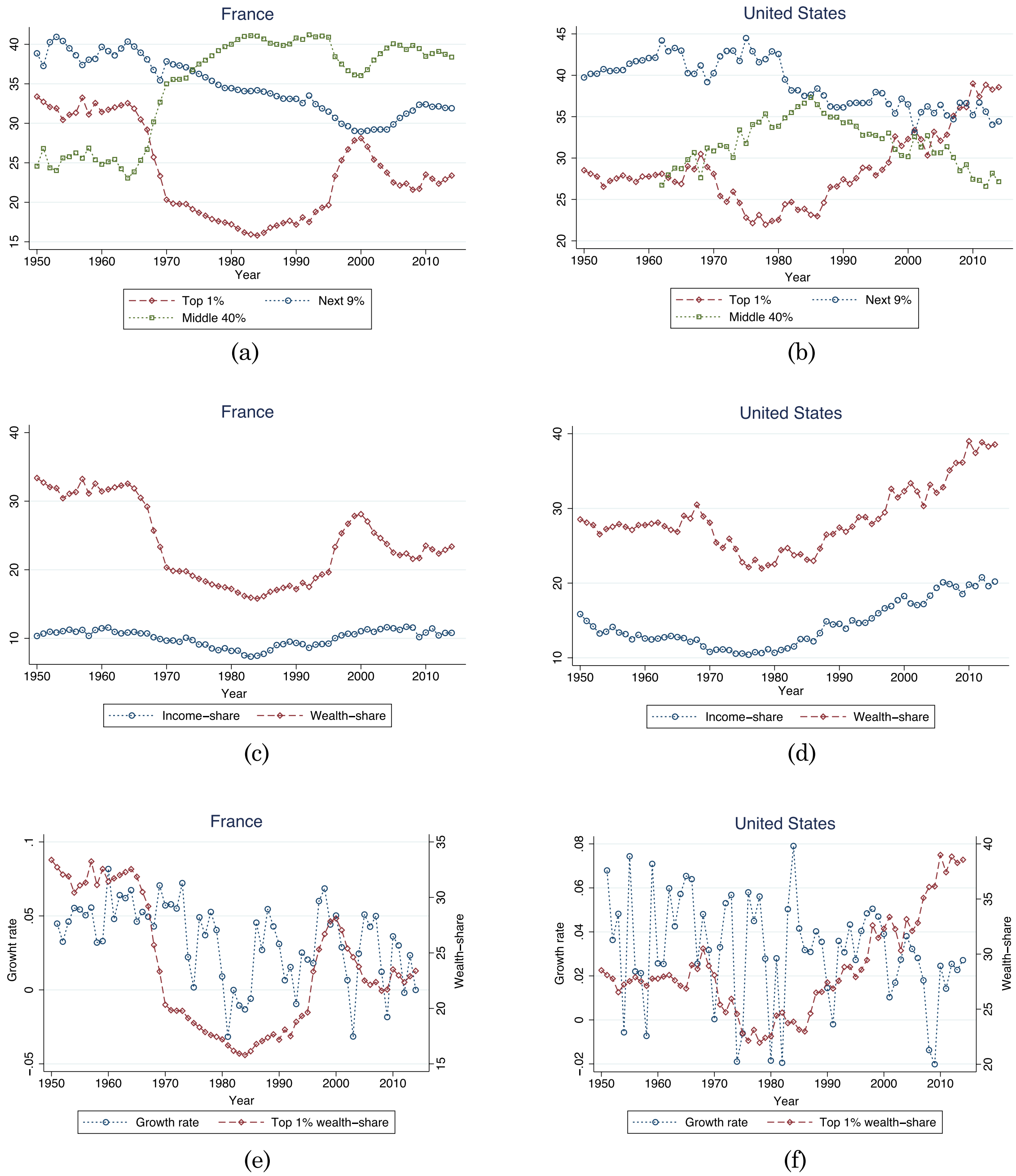

Fig. 1. The top 1% in perspective.

review, see Herzer and Vollmer, 2012; Berg et al., 2018). Nonetheless, the results have not been conclusive, and are even frequently contradictory.

For instance, among recent studies, Castelló-Climent (2010) report that although higher income inequality reduces growth in low- and middle-income countries, it has a zero or positive effect in high-income ones. Conversely, Brueckner and Lederman (2018) show that in low-income countries, growth is boosted by greater income inequality, whereas in high-income nations, inequality has a significant negative effect on transitional growth. Beyond differences in the quality of the data and the econometric technique used, the empirical difficulty to identify a monotonic relationship reflects the heterogeneity underlying theoretical mechanisms, as has been previously discussed. For example, Halter et al. (2014) find that a deterioration of income distribution helps economic performance in the short-term but reduces the output growth rate in the long-term. The studies by Chambers and Krause (2010) and El-Shagi and Shao (2019) indicate that inequality is either beneficial or detrimental to growth depending on income and educational conditions. Notably, in a disaggregated set-up, Erman and te Kaat (2019) establish that an unequal income distribution increases the growth rate of physical capital-intensive industries and reduces the growth rate of human capital-intensive ones. Furthermore, Litschig and Lombardi (2019) present evidence that, in terms of the subsequent growth effects, it matters whether inequality or redistribution originates from the lower or upper tails of the distribution.

Specifically, regarding top income, Herwartz and Walle (2020) have showed that an increase in the share of income that goes to the top 1 percent is related to higher economic activity. Focusing on 12 OECD economies in the post-1950 period, the main transmission channel lies in the prevailing structure of incentives. Nonetheless, considering the top 1 percent wealth share, Islam and McGillivray (2020) find a negative correlation based on a panel for 45 countries. However, the implications of their study are rather limited by the short period of analysis, i.e., 2000–2012. Bagchi and Svejnar (2015) also obtain a negative correspondence upon building a *proxy* for wealth distribution based on Forbes list of billionaires. They generate three measures of wealth inequality by dividing the sum of all billionaires' wealth by either output, the capital stock, or population. Our understanding is that these indicators are not wholly intuitive and a more comprehensive approach is still needed.

The last three studies revisited herein are very much related to our contribution as we are also interested in those who are the very top of the distribution. Nonetheless, contrasting with the first, our focus lies on the growth–*wealth* share nexus. Moreover, compared to the other two, our time dimension is sufficiently large to arrive at more robust conclusions. We also go one step further and provide some insights on the role of the top 0.01 percent.

3. The top 1 percent in perspective

There are many reasons why societies might be concerned about inequality and desire to measure it. Thus, the selection of the index to measure income or wealth distribution becomes highly important. The choice of such an index embodies fundamental normative judgements that are important to be considered when interpreting any results (for a recent review of different indicators, see Nino-Zarazúa et al., 2017). As indicated by Alvaredo et al. (2013), in recent years, there has been a growing realisation that rising inequality in the US and Europe is very much related to what has been happening at the very top. Hence, we should start by looking at the factual importance of the top 1 percent. On the one hand, data from the *World Inequality Database* indicate that, in 2014, the average net personal wealth of this group is US$ 5.8 million in France and approximately US$ 12 million in the US, while its average pre-tax income, one the other hand, is US$ 500 thousand in the former country and US$ 1.4 million in the latter.[1]

Using the same source, Fig. 1 (a) and (b) presents the time series of wealth distribution of the top 1 percent, the next 9 percent, and the middle 40 percent in these two countries. While until the 1980s, the wealth share going to the first group had been stable or decreasing, this trajectory was reverted afterwards. In France, it fell from 30% to 15% and seemed to have stabilised close to 25%. In the US, however, the top 1 percent wealth share fell from 30% to 20% in 1980 but had consistently grown since then, almost reaching the 40% threshold in 2014. These trajectories are in sharp contrast with the wealth share controlled by the other two groups. For instance, on the one hand, the French next 9 percent has presented an approximately monotonic reduction from 40% to 30%. The middle 40 percent, on the other hand, was stable from 1950 to 1970, around 25%, and jumped to 40% during the seventies, apparently stabilising afterwards. Conversely, both groups in the US have experienced a reduction in their share over the past three decades.

Referring to France, Garbinti et al. (2020) show that rising inequality in saving rates coupled with highly stratified rates of return has led to higher wealth concentration in spite of the opposing effect of house price increases. They also indicate the effect of asset price movements on wealth share fluctuations. A similar conclusion is reached by Saez and Zucman (2016) for the US, suggesting that besides the saving rate effect, the increase in wealth inequality in recent decades is due to the upsurge of top income (also see Mollick, 2012). These dynamics further motivate our interest in the top 1 percent.

We can also appreciate the extent to which wealth inequality might be more relevant than income inequality. Davies et al. (2011), for example, find that global wealth-holding is by far more concentrated than income. While in most countries the Gini coefficient for disposable income lies within the range of 0.3–0.5, regarding wealth inequality, it is normally between 0.6 and 0.8. Fig. 1 (c) and (d) shows that both dimensions of inequality move together, but wealth is two to three times more concentrated. In 2014, the US top 1 percent controlled 20% of the income and approximately 40% of the total wealth.

At least in the US, rising inequality has generated a substantial increase in savings by the top and debt by the bottom of the income distribution that has not been associated with an increase in investment (e.g. Cynamon and Fazzari, 2016; Mian et al., 2020). Fig. 1 (e) and (f) shows plots, in red, of the top 1 percent wealth share and, in blue, of the output growth rate. The first thirty years of our sample are characterised by higher and more volatile growth with stably decreasing wealth concentration. However, the last twenty to thirty years are marked by relatively lower and less volatile growth, with an increasing share of wealth going to the top 1 percent. Such trajectories might suggest a negative correspondence between our relevant variables. For instance, in the US, the period referred to as the Great Moderation was marked by increasing wealth concentration. However, as indicated in the previous Sections, we have reasons to believe that this relationship is not monotonic. Hence, we shall proceed by investigating the properties of this correlation.

4. The inverted-U

A common set-up for investigating the relationship between the output growth rate (g) and a vector of explanatory variables (X) is

[1] The *World Inequality Database* was initially created as the *World Top Incomes Database*, providing access to all the existing series constructed by T. Piketty, E. Saez, and co-authors in past years. Its key novelty was the combination of fiscal, survey, and national accounts data in a systematic manner. This allowed the computation of more reliable top income shares series than previous inequality sources.

as follows:

$$g_t = \psi y_{t-1} + \gamma X_{t-1} + \varepsilon_t \tag{1}$$

where y is the logarithmic of the gross domestic product (GDP); ψ and γ are the coefficients associated with y and X; and ε represents the error term. Recalling that $g_t \approx y_t - y_{t-1}$, this is equivalent to estimating:

$$y_t = (1 + \psi)y_{t-1} + \gamma X_{t-1} + \varepsilon_t \tag{2}$$

The use of time-averaged data in a panel context is common practice in the inequality–growth literature. This approach allows the inclusion of a larger number of economies while it is supposed to eliminate business cycle effects. However, it has several shortcomings. One of the main criticisms is the implicit assumption of a common economic structure across countries. Panel methods permit controlling for country-specific omitted variables. However, in dynamic models, they might produce inconsistent and potentially misleading estimates when the slope of coefficients differs across cross-sectional units (e.g. Pesaran and Smith, 1995; Attanasio et al., 2000). Moreover, averaging data over time may induce spurious contemporaneous correlations. Both their sign and magnitude might differ from those in the underlying data, a problem not to be solved by instrumental variable estimation, nor even the generalised method of moments (see Ericsson et al., 2001). Finally, annual data provides information that is lost when averaging, especially when series are highly persistent (Nair-Reichert and Weinhold, 2001). It is unclear if business cycle effects are effectively eliminated as the length of the interval over which averages are computed is arbitrary (Herzer and Vollmer, 2012).

Therefore, we shall subsequently rely on time series estimation techniques. Our aim is to study the correspondence between g and the top 1 percent wealth share (ω) in France and the US, while considering the possibility of non-linearities in the underlying long-run relationship. Therefore, we utilise different specifications of Eqs. (1) and (2), including alternative controls in the vector of explanatory variables. In Section 5, we proceed by performing a series of robustness checks. We shall also investigate the existence of such a correspondence for the top 0.01 percent.

4.1. Dataset and order of integration

Our dataset is annual and covers the period between 1950 and 2014. The output growth rate is obtained as the difference of the logs of the output-side real GDP at chained purchasing power parity from the *Penn World Table* (PWT) version 9.0. The top 1 percent wealth shares are obtained from the *World Inequality Database*. Our set of controls comes from the PWT and includes the size of the population (Pop), the capital stock, a human capital index (HC), the share of government consumption on GDP (Gov), and the degree of trade openness. This last variable is measured as the sum of exports and imports over GDP. We also include dummy variables for each decade to consider the possibility of structural breaks. They assume a value of 1 for years belonging to the respective decade, and zero otherwise.

As mentioned in the Introduction, our choice of working with the French and US economies is basically due to data availability. Both countries also allow us to virtually represent alternative institutional arrangements. Notwithstanding, one should recognise the existence of a significant level of heterogeneity within continental Europe. A remarkably more equal welfare regime is missing, namely, the Nordic countries. While also displaying a high level of wealth concentration, this last group is less wealth-unequal than both France and the US. For instance, Ranaldi and Milanovic (2021) have recently shown that Finland, Norway, Denmark, and Iceland are quite exceptional because they have high compositional inequality, i.e. how the composition of income between capital and labour varies along income distribution, with low interpersonal inequality (see also Iacono and Ranaldi, 2020). Future research covering those experiences is to be encouraged once data become available.

Ascertaining the order of integration of the variables under analysis is an essential precondition for choosing a consistent estimator. Consequently, we perform the augmented Dickey–Fuller (ADF) test, with and without a break, and the Phillips–Perron (PP) test. The results are reported in the Empirical Appendix, and they indicate that g is stationary, I(0), while y, ω, and ω^2 are integrated of order one, I(1). Hence, we use the autoregressive distributed lag (ARDL) cointegrating estimator, developed by Pesaran and Shin (1998) and Pesaran et al. (2001). This methodology has the advantage of allowing analysis regardless of whether variables are a mixture of I(0) or I (1), which is our case.

Even though we remain constrained by parametric assumptions, our approach overcomes two of the main shortcomings of usual cross-country regressions. On the one hand, by focusing exclusively on the time dimension of the data, we avoid the well-known heterogeneity problems of a panel set-up. On the other hand, under cointegration, the omitted variable issue is unlikely to affect the reliability of our estimates. This is because an omitted variable will either be stationary – in which case the estimated coefficients are invariant to its inclusion – or non-stationary – in which case we will not be able to obtain a stable cointegrating relationship if we exclude it.[2]

[2] The presence of cointegration brings a form of robustness to many of the classic empirical problems that lead to the violation of the so-called exogeneity condition for the regressors. Examples include measurement error, simultaneity, omitted variables, reverse causality, or anything that leads the data generating process. For a review of the super-consistency properties under cointegration, see Pedroni (2019).

4.2. Estimation and main results

Based on Eq. (1), our general ARDL(p, q, l) model is given by:

$$g_t = \alpha_0 + \alpha_1 t + \sum_{i=1}^{p} \psi_i g_{t-i} + \sum_{j=0}^{q} \beta_{\omega,j}\omega_{t-1-j} + \sum_{k=0}^{l} \beta_{\omega^2,k}\omega_{t-1-k}^2 + \gamma X_{t-1} + \varepsilon_t \tag{3}$$

while in terms of Eq. (2), we have:

$$y_t = \alpha_0 + \alpha_1 t + \sum_{i=1}^{p} \psi_i y_{t-i} + \sum_{j=0}^{q} \beta_{\omega,j}\omega_{t-1-j} + \sum_{k=0}^{l} \beta_{\omega^2,k}\omega_{t-1-k}^2 + \gamma X_{t-1} + \varepsilon_t \tag{4}$$

where α_0 is a constant term; α_1 and ψ_i are the coefficients associated with a linear trend and lags of g or y, respectively; while $\beta_{\omega,j}$ and $\beta_{\omega^2,k}$ represent the lag coefficients of ω and ω^2, respectively.[3]

In an infinite distributed lag model, an infinite number of lag weights need to be estimated. When the samples are finite, however, we have to introduce an assumption about the maximum number of lags beyond which values of the independent variable do not affect the dependent one. The Akaike information criterion (AIC) provides a rule for model selection dealing with the trade-off between goodness of fit and simplicity. For our objectives herein, we thoroughly rely on the AIC, imposing a maximum of four lags for dependent and independent variables. This means that, for each set of controls, we estimate 100 times the model with different lag combinations and select the one that performs best according to the AIC.[4]

If two series are cointegrated, this means that they have a long-term relationship which prevents them from wandering apart without bound. Pesaran et al. (2001) and Narayan (2005) provide supply bounds on the critical values for the asymptotic distribution of the F- and t-statistics. Tables 1 and 2 present our main results for France and the US, respectively. We identify a statistically significant quadratic long-run relationship between our relevant variables. The estimated function is robust to the inclusion of different exogenous controls. To simplify our presentation, we have omitted the short-term coefficients, which are available under request.

Regarding models I–IV, in France, the coefficient of the linear term varies between 0.021 and 0.044, while the quadratic term is in the -0.0004 to -0.0009 interval. However, in the US, the quadratic coefficient has a similar magnitude to its European counterpart. Nonetheless, the linear portion of the relationship is slightly higher, lying between 0.025 and 0.05. We test for predictive causality in the sense of Granger by incorporating the lagged error-correction (EC) term that represents the long-run causal relationship. In all cases, a negative and significant EC coefficient implies that there is convergence to the long-run equilibrium solution. Models I and II show an EC $\in [-2, -1]$, which means they produce dampened fluctuations towards equilibrium (Loayza and Rancière, 2006). Once some further controls are included, we obtain an EC $\in [-1, 0]$. Movements into disequilibrium between 70% and 90% are corrected for within one period.

As for models VI–X, the coefficients are slightly higher because now we are estimating the impact of wealth concentration on the level of output. However, the quadratic correspondence remains. In France, the linear part of the relationship lies between 0.10 and 0.17, while the quadratic term is in between -0.002 and -0.003. Upon adding exogenous controls, the EC term increases in magnitude and stabilises such that 25–30% of any deviation from the equilibrium solution is corrected for within one period. The existence of a long-run relationship is confirmed by the F- and t-statistics. As evident in models VII–IX, the 5% significance threshold is overcome. For the US, once we control for the stock of physical and human capital, we recover our cointegrating vector. Nevertheless, it is noticeable that upon further including government consumption and the degree of trade openness, our coefficients lose statistical significance. The linear part of the relationship lies between 0.08 and 0.21, while the quadratic term is in between -0.0015 and -0.0039.

Given that our estimates suggest that the output growth rate is an inverted-U-shaped function of the wealth share of the top 1 percent, it is possible to compute the local *optimal wealth share* as the share of wealth which is compatible with the maximum growth rate. Divergences from this level in any direction are associated with a reduction in long-run growth. The last line of Tables 1 and 2 indicates that such a local optimal is approximately 25%, being slightly higher in the US than in France before controlling for the size of the government and trade openness. Notably, this level of wealth concentration is not a sufficient condition for maximal economic growth; we only claim that ω^* is compatible with it. Fig. 2 depicts the *Kuznets curve of the Rich* and the derivative of g with respect to ω for models III–V. The estimated "optimal" corresponds to the point where the zero-dashed line is crossed. For the values of $\omega < \omega^*$, g is an increasing function of the wealth share. Once $\omega > \omega^*$, the partial derivative becomes negative, and further increases in wealth concentration are associated with a reduction in economic performance.

Economics has many quadratic relationships. Lind and Mehlum (2010) provide the necessary and sufficient conditions for correctly testing a U-shape in finite samples. In the context of this paper, given that $\omega \in [0, 100]$, we need to have:

[3] In the case of an ARDL (1,0,0), for example, the correspondent estimated equations are $g_t = \alpha_0 + \alpha_1 t + \psi_1 g_{t-1} + \beta_{\omega,0}\omega_{t-1} + \beta_{\omega^2,0}\omega_{t-1}^2 + \gamma X_{t-1} + \varepsilon_t$ and $y_t = \alpha_0 + \alpha_1 t + \psi_1 y_{t-1} + \beta_{\omega,0}\omega_{t-1} + \beta_{\omega^2,0}\omega_{t-1}^2 + \gamma X_{t-1} + \varepsilon_t$. The reported long-run coefficients result from extensive manipulations of these expressions. For a specialised reader interested in the econometric details, please see Pesaran and Shin (1998) and Pesaran et al. (2001).

[4] Even though AIC over-specifies the model for all values of the true lags, we prefer it over the Schwarz information criterion because it is more efficient. Whenever serial correlation was found, we adjusted the number of lags accordingly.

Table 1
Growth and the 1% top wealth share – France.

	Dependent variable: g_t					Dependent variable: y_t				
cline2-11	I	II	III	IV	V	VI	VII	VIII	IX	X
	ARDL(4,1,0)	ARDL(4,1,0)	ARDL(1,0,1)	ARDL(1,0,1)	ARDL(1,0,0)	ARDL(2,2,0)	ARDL(2,1,0)	ARDL(2,0,1)	ARDL(2,0,1)	ARDL(1,0,2)
ω_{t-1}	0.026661***	0.021230***	0.042191***	0.044655***	0.023436***	0.670118	0.173786***	0.107486***	0.128290**	0.122416***
ω^2_{t-1}	− 0.000534***	− 0.000429***	− 0.000894***	− 0.000948***	− 0.000412***	− 0.013850	− 0.003703***	− 0.002278***	− 0.002725**	− 0.002503**
y_{t-1}	–	–	✓	✓	✓	–	–	–	–	–
Pop_{t-1}	–	–	✓	✓	✓	–	–	✓	✓	✓
Capital stock_{t-1}	–	–	–	✓	✓	–	–	–	✓	✓
HC_{t-1}	–	–	–	✓	✓	–	–	–	✓	✓
Gov_{t-1}	–	–	–	–	✓	–	–	–	–	✓
Openess_{t-1}	–	–	–	–	✓	–	–	–	–	✓
Time dummies	–	✓	✓	✓	✓	–	✓	✓	✓	✓
EC	− 1.183436***	− 1.383853***	− 0.743183***	− 0.741968***	− 0.838851***	− 0.033058***	− 0.160156***	− 0.291721***	− 0.258262***	− 0.301949***
Bounds F-stat.	11.89728	13.57796	16.76351	16.70782	11.84733	3.474326	7.362544	8.943439	6.541070	6.149311
Bounds t-stat.	− 6.090281	− 6.522596	− 7.234860	− 7.228857	− 5.754013	− 3.287706	− 4.792821	− 5.284452	− 4.523079	− 3.1583
LM F-stat	0.087211	0.494021	1.623887	1.728071	0.750015	0.268345	1.145680	1.623887	1.728071	0.058073
ω^*	24.9634	24.7435	23.5967	23.5522	28.4417	24.1919	23.4655	23.5921	23.5394	24.4538

*, **, *** stand as 10%, 5%, and 1% of significance.

Table 2
Growth and the 1% top wealth share – United States.

	Dependent variable: g_t					Dependent variable: y_t				
	I	II	III	IV	V	VI	VII	VIII	IX	X
	ARDL(1,2,0)	ARDL(3,2,1)	ARDL(1,2,1)	ARDL(1,1,1)	ARDL(1,1,1)	ARDL(2,2,0)	ARDL(1,2,1)	ARDL(2,2,1)	ARDL(1,1,1)	ARDL(1,1,1)
ω_{t-1}	0.008177	0.025966**	0.052312***	0.049682**	0.049422*	0.083264***	0.210810**	0.150450**	0.126448**	0.121572
ω^2_{t-1}	− 0.000143	− 0.000482**	− 0.000972***	− 0.000937**	− 0.000945*	− 0.001567***	− 0.003908*	− 0.002795**	− 0.002409**	− 0.002346
y_{t-1}	–	–	✓	✓	✓	–	–	–	–	–
Pop_{t-1}	–	–	✓	✓	✓	–	–	✓	✓	✓
$\text{Capital stock}_{t-1}$	–	–	–	✓	✓	–	–	–	✓	✓
HC_{t-1}	–	–	–	✓	✓	–	–	–	✓	✓
Gov_{t-1}	–	–	–	–	✓	–	–	–	–	✓
Openess_{t-1}	–	–	–	–	✓	–	–	–	–	✓
Time dummies	–	✓	✓	✓	✓	–	✓	✓	✓	✓
EC	− 0.909584***	− 1.426662***	− 0.847356***	− 0.933362***	− 0.901379***	− 0.202737***	− 0.217234***	− 0.294627***	− 0.355576***	− 0.355028***
Bounds *F*-stat.	15.46214	15.44247	13.43589	19.38177	14.59464	2.798331	3.061499	3.463931	6.948717	6.151998
Bounds *t*-stat.	− 6.933484	− 6.956034	− 5.825986	− 7.789312	− 6.765656	− 2.950579	− 3.093087	− 2.527100	− 4.661890	− 3.341079
LM *F*-stat.	0.062140	0.609969	0.982702	1.522882	1.455087	0.442009	0.027167	0.982702	0.983259	0.862077
ω^*	28.5909	26.9356	26.9094	26.5112	26.1492	26.5679	26.9715	26.9141	26.2449	25.9104

*, **, *** stand as 10%, 5%, and 1% of significance.

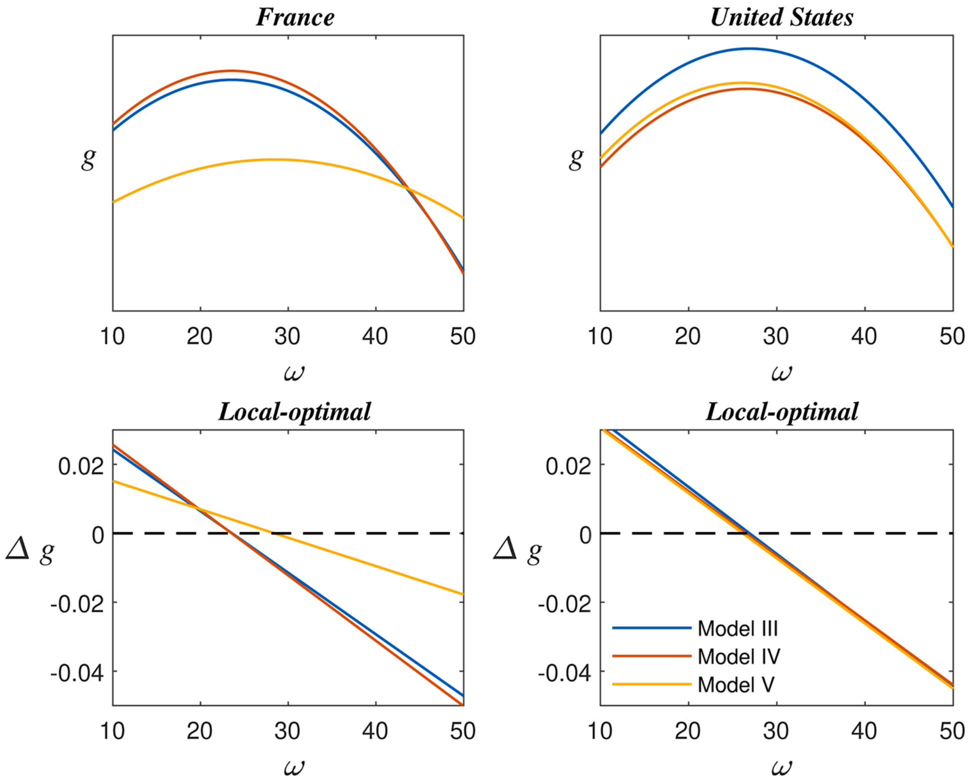

Fig. 2. Kuznets curve of the rich and the determination of the local optimal.

$$\frac{\partial g}{\partial \omega}\Big|_{\omega=0} > 0$$

and

$$\frac{\partial g}{\partial \omega}\Big|_{\omega=100} < 0$$

Through direct computation, it is easy to see that the two conditions are satisfied for both countries, thereby complementing the visual representation in Fig. 2.

Our findings have some similarities with Banerjee and Duflo (2003). Using non-parametric and cross-country techniques, they demonstrate that the rate of growth is a function of net changes in inequality such that variations in inequality, in any direction, are associated with reduced growth in the next period. We can also directly compare our estimates to Islam and McGillivray (2020) who report that an increase in the wealth share of the top 1 percent is associated with a reduction in growth. This happens in our case only after the local optimal value. Given that their study used post-2000 data, when wealth was already highly concentrated, this could indicate that only the downward part of the curve was effectively captured. Hence, their findings indicating that wealth inequality is detrimental for growth is not particularly surprising, and they appear to be consistent with our results.

In the previous section, we show that the wealth share of the top 1 percent in France is slightly below the 25% threshold. However, the wealth controlled by the 1 percent elite in the US has recently approached 40%. Based on Fig. 2, this means that while the former country is closer to its long-run growth potential, the latter is growing between 1% and 2% less. The overconcentration of wealth in the hands of few has been beleaguering the economic performance of the US, especially after the 2000s.

The main underlying transmission channel lies in the prevailing structure of incentives. On the one hand, very low wealth shares might indicate insufficient skill premia. If agents with a particular set of skills – ranging from top performers to corporate officers and decision makers – are not adequately rewarded, there may be a mismatch between the distribution of resources and social costs – benefits. This could lead to a system with an inefficient allocation of resources. The so-called trickle-down effect postulates that economic growth generated by the rich eventually benefits the poor through job creation and other opportunities. Comparatively, our estimates may support this view. However, high wealth concentration is likely to induce the consolidation of equally inefficient institutions in public and private spheres. This includes the obstruction of the access of people with new ideas and skills to the economy, i.e., limiting the opportunities of those who are not already at the top. In the long-term, there is a subversion of institutions, resulting in an ineffective political system and dysfunctional markets.[5]

Moreover, while inequality might be positively connected to incentive rewards, wealth accumulation has an additional intergenerational component that functions in the opposite direction. Alvaredo et al. (2017) have provided historical series on the evolution of the share of inherited wealth in aggregate wealth for different European countries and the US. Inherited wealth represented 80–90% of the total wealth in France in the 19th century, and this share fell to 40–50% during the 20th century. However, it is back to approximately 60–70% in the early 21st century. Current rates in the US are not very different, with potential implications for economic growth. By maintaining the notion that incentives matter, we highlight how detrimental an overconcentration of wealth might be in terms of growth. This is consistent with recent evidence indicating that rising inequality in the US has increased saving rates with no impact on capital accumulation (see Mian et al., 2020).

Redistributive issues have a political economy of their own, and it is not our intention to be irresponsible in those matters. Notably also, the success of any policy implementation is considerably conditional to its design. In a recent and detailed study on the topic, Berg et al. (2018) have shown that redistribution appears benign in terms of its impact on growth, except when it is extensive. Saez and Zucman (2019) have advocated for an “economic income” tax schedule calculated on the basis of wealth. Further research on the impacts of wealth redistribution is certainly to be encouraged.

5. Exploring the robustness of the inverted-U

By estimating an ARDL version of two alternative specifications of the growth equation, we are able to identify an inverted-U relationship between wealth concentration at the top 1 percent and growth. In this Section, we aim to check the robustness of such a result. This is done in three different ways: first, we adopt the flexible Nonlinear ARDL (NARDL) estimator proposed by Shin et al. (2014) that allows us to simultaneously and coherently model asymmetries both in the cointegrating vector and in the patterns of dynamic adjustment. Second, given that Eq. (2) confronts series that are non-stationary and integrated of order one, we utilise the dynamic ordinary least squares (DOLS) estimator developed by Stock and Watson (1993). This estimator addresses the problem of second-order asymptotic bias arising from serial correlation and endogeneity. Finally, we also check if a similar correspondence can be verified for the top 0.01 percent. Considering that some authors, such as Garbinti et al. (2020), have shown that asset price movements have an important impact on wealth share variations, DOLS regressions further control for the rates of return on equity (EqR), housing (HousR), and government bonds (GovBR), as obtained from Jordà et al. (2019).[6]

5.1. Allowing for asymmetric effects

The nonlinearity of many macroeconomic variables and processes has long been recognised in the economic literature. Because asymmetric effects are fundamental to the human condition, we abandon the assumption that the long-run relationship may be represented as a symmetric linear combination of regressors. Therefore, the Nonlinear ARDL estimator provides a straightforward means of testing both long- and short-run symmetry restrictions.

Let $\omega = \omega_0 + \omega^+ + \omega^-$ and likewise $\omega^2 = \omega_0^2 + \omega^{2+} + \omega^{2-}$. The superscripts $+$ and $-$ indicate the partial sum processes of positive and negative changes in the wealth share:

$$\omega_t^+ = \sum_{i=1}^{t} max\{\Delta\omega_i, 0\} \omega_t^- = \sum_{i=1}^{t} min\{\Delta\omega_i, 0\}$$

$$\omega_t^{2+} = \sum_{i=1}^{t} max\{\Delta\omega_i^2, 0\} \omega_t^{2-} = \sum_{i=1}^{t} min\{\Delta\omega_i^2, 0\}$$

We consider the following NARDL (p, q, l) model:

[5] As pointed out by one of the reviewers, skills are non-rival and there are important complementarities between them such that the society as a whole benefits from more – as opposed to fewer – people being skilled. Nevertheless, skills are not a pure public good because people can be prevented from access to education and training, excluded from certain social networks, etc. Hence, they are not equally distributed among the population.

[6] The DOLS estimator is preferred to single equation alternatives, such as the fully modified ordinary least squares (FMOLS), because it enables the differentiation of leads and lags when specifying the model. This has proven to be useful for controlling serial correlation problems. Moreover, Kao and Chiang (2001) have shown, in a panel context, that the DOLS estimator outperforms FMOLS regressions as the former is computationally simpler.

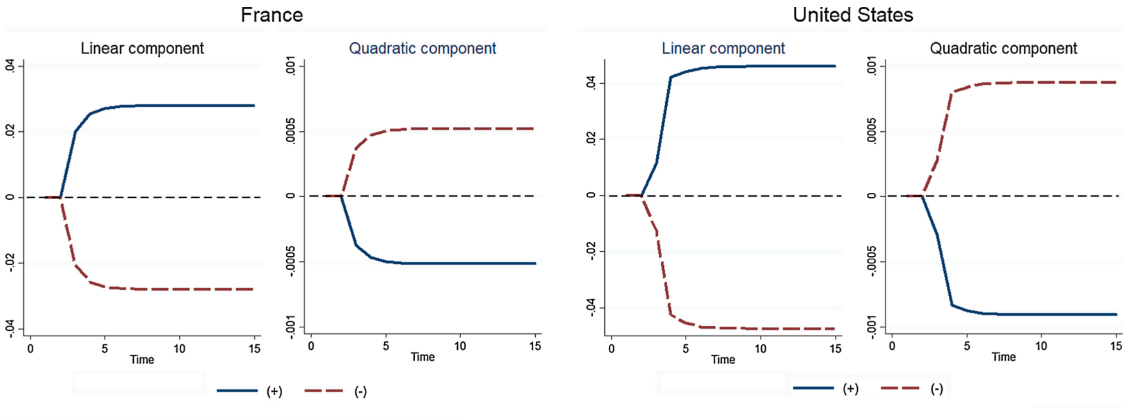

Fig. 3. Convergence to the long-run solution in an NARDL model.

Table 3
DOLS estimations - France.

Dependent variable: GDP_t	XI	XII	XIII	XIV	XV
ω_{t-1}	0.122176***	0.025298**	0.028325**	0.032468**	0.056036***
ω^2_{t-1}	− 0.002789***	− 0.000527**	− 0.000574**	− 0.000647**	− 0.001115**
α_0	11.67638***	− 0.097293	− 0.862095	0.844950	0.894709
α_1	0.026797***	0.002859	0.000419	0.004443	0.005671
GDP_{t-1}	–	0.679245***	0.668748***	0.508176***	0.364813***
$Pop._{t-1}$	–	1.047613	1.211097	0.931625	1.160164*
Capital $stock_{t-1}$	–	–	0.009811	0.169825*	0.234693**
HC_{t-1}	–	–	0.541522	2.530552***	2.935162***
Gov	–	–	–	− 1.645221***	− 0.977771*
$Openess_{t-1}$	–	–	–	− 0.003918***	− 0.004377***
EqR	–	–	–	–	0.010282
HousR	–	–	–	–	0.090178
GovBR	–	–	–	–	− 0.059097**
Time dummies	✓	✓	✓	✓	✓
Lc stat.	0.060489	0.132536	0.154888	0.185702	0.222185
Jarque Bera	2.704162	2.124455	2.809664	0.140589	0.162936
ω^*	21.9031	24.0018	24.6733	25.0911	25.1282

*, **, *** stand as 10%, 5%, and 1% of significance.

$$g_t = \alpha_0 + \alpha_1 t + \sum_{i=1}^{p} \psi_i g_{t-i} + \sum_{j=0}^{q} (\beta_j^+ \omega_{t-1-j}^+ + \beta_j^- \omega_{t-1-j}^-) + \sum_{k=0}^{l} (\beta_k^+ \omega_{t-1-k}^{2+} + \beta_k^- \omega_{t-1-k}^{2-}) + \gamma X_{t-1} + \varepsilon_t \quad (5)$$

Fig. 3 shows trajectories converging to the long-run solution when we re-estimate model V, which contains all our controls. A positive shock in the wealth share of the top 1 percent corresponds to the blue line, while the red dashed one represents a negative shock. We are unable to identify significant asymmetries in the underlying long-run relationship. Notwithstanding, given that the linear and quadratic components have opposite signs, it is possible to appreciate the inverted-U. Moreover, in both countries, the adjustment to equilibrium takes approximately five years to occur, being slightly faster in the US.

5.2. Dynamic OLS

Using Eq. (2) has the advantage of confronting variables that are integrated of order one. Hence, we estimate the following DOLS model:

Table 4
DOLS estimations – United States.

Dependent variable: GDP_t	XI	XII	XIII	XIV	XV
ω_{t-1}	0.099442**	0.066128***	0.055463***	0.043161**	0.030096**
ω^2_{t-1}	− 0.001752**	− 0.001200***	− 0.001008***	− 0.000833***	− 0.000526**
α_0	13.23346***	4.092829*	11.06586***	10.88420***	9.319126***
α_1	0.037625***	0.015074***	0.023582***	0.024055***	0.019627***
GDP_{t-1}	–	0.686347***	0.587028***	0.644972***	0.631123***
$Pop._{t-1}$	–	− 0.079506	− 0.141938	− 0.110494	− 0.028356
Capital $stock_{t-1}$	–	–	− 0.885167***	− 0.952723***	− 0.734565***
HC_{t-1}	–	–	0.391599***	0.318550***	0.328046***
Gov_{t-1}	–	–	–	0.195369	− 0.086132
$Openess_{t-1}$	–	–	–	− 0.000137	− 0.001341
EqR_{t-1}	–	–	–	–	0.074507***
$HousR_{t-1}$	–	–	–	–	0.031829
$GovBR_{t-1}$	–	–	–	–	0.007033
Time dummies	✓	✓	✓	✓	✓
Lc stat.	0.056466	0.130968	0.232933	0.263222	0.353333
Jarque Bera	0.620701	0.215894	1.046201	1.701618	1.117368
ω^*	28.3795	27.5533	27.5114	25.9069	28.6083

*, **, *** stand as 10%, 5%, and 1% of significance.

Table 5
Looking at the top 0.01 percent.

Dependent variable: GDP_t	France			United States		
	XVI	XVII	XVIIII	XIX	XX	XXI
ω_{t-1}	0.072467*	0.041618	0.031524	0.053668***	0.067763***	0.033655***
ω^2_{t-1}	− 0.008740	− 0.003147	− 0.001817	− 0.004771***	− 0.005647***	− 0.002457***
α_0	0.024522	1.143977	0.438641	12.03162***	9.890173***	10.61049***
α_1	0.002234	0.004848	0.004801	0.023396***	0.018903***	0.020642***
GDP_{t-1}	0.771964***	0.702605***	0.668375***	0.504069***	0.456898***	0.530729***
$Pop._{t-1}$	0.727374	0.401617	0.679173	− 0.242404	− 0.051354	− 0.227150
Capital $stock_{t-1}$	− 0.013823	0.118531	0.157496	− 0.681018***	− 0.385144	− 0.504611***
HC_{t-1}	0.336568	2.042518***	1.358944**	0.400441***	0.362141***	0.340713***
Gov_{t-1}	–	− 2.086683***	− 1.519653***	–	− 0.603165	− 0.466821
$Openess_{t-1}$	–	− 0.002727***	− 0.002480**	–	− 0.001140	− 0.001943
EqR_{t-1}	–	–	0.022678***	–	–	0.071363***
$HousR_{t-1}$	–	–	0.067253*	–	–	0.054916
$GovBR_{t-1}$	–	–	− 0.027948	–	–	0.006329
Time dummies	✓	✓	✓	✓	✓	✓
Lc stat.	0.150725	0.237741	0.275422	0.238601	0.206387	0.417335
Jarque Bera	0.574080	0.958050	0.799523	0.715714	0.779969	0.572519
ω^*	4.1457	6.6123	8.6747	5.6243	5.9999	6.8487

*, **, *** stand as 10%, 5%, and 1% of significance.

$$y_t = \alpha_0 + \alpha_1 t + \sum_{j=-q}^{q} \beta_{\omega,j} \Delta\omega_{t-1-j} + \sum_{k=-l}^{l} \beta_{\omega^2,k} \Delta\omega^2_{t-1-k} + \gamma X_{t-1} + \varepsilon_t \tag{6}$$

As evident in Table 3, our previous results are basically confirmed for France. Model XI does not include lagged GDP as an exogenous control. In this case, the obtained coefficients are similar to those in the last five columns of Table 1. However, models XII–XV do include lagged GDP as an explanatory variable, leading to estimates closer to what we reported when estimating Eq. (1). Model XV also controls for the rates of return on equity, housing, and government bonds. Coefficients of the linear part are in between 0.025 and 0.04. For the quadratic term, we have a negative effect of −0.0005 to −0.0008. Hansen cointegration tests provide a sufficiently low Lc statistic, so we cannot reject the null hypothesis that the series are cointegrated. A local *optimal wealth share* close to 25% also agrees with our previous findings.[7]

For the US, DOLS estimations appear to be more robust than the ARDL alternative. In model XI, on Table 4, the linear part is slightly lower than in France, being equal to 0.099. The quadratic coefficient also has a smaller magnitude, being equal to −0.001. The remaining models have coefficients of the linear part lying in between 0.04 and 0.06, while the quadratic component oscillates from −0.0008 to −0.0012. In all specifications, we cannot reject the null of cointegration and estimated parameters are significant. Furthermore, the local *optimal wealth share* of the top 1 percent is slightly above the 25% line. As a robustness check, in the Empirical Appendix, we further investigate whether the relationship holds if we use Net National Income (NNI), a macroeconomic aggregate commonly employed in the inequality literature, instead of GDP. We show that this is the case for both countries.

Given the emphasis of this article on the share of wealth controlled by the upper class, we present a final set of tests for the top 0.01 percent of the population in Table 5. In 2014, this group had an average net personal wealth of US$ 70 million in France and approximately US$ 350 million in the US. The inverted-U function is robust with a linear coefficient ranging in between 0.03 and 0.07, and the quadratic term lying in the interval of −0.002 to −0.008. However, in the case of France, our estimates are no longer statistically significant, implying that the inverted-U weakens as we move from the rich to the super-rich. In the US, on the other hand, our results are significant at 1%, leading to a local optimal between 5.5% and 7%. Recalling that since 2010 the net personal wealth share of this group has been above 10%, this is remarkably higher than what our estimations suggest to be "optimal". Such a result confirms one of the main insights of the article, i.e., the overconcentration of wealth in the hands of few has damaged economic performance in the US. Moreover, it complements the findings of Mian et al. (2020) that show that rising inequality in the US has been associated with a large increase in savings by the top of the income distribution, higher debt at the bottom, and no effect on investment. It is also in line with Saez and Zucman (2016) who demonstrate that increments in the share of wealth owned by the top 1 percent are driven by the top 0.01 percent.

Overall, we provide robust evidence that the output growth rate is an inverted-U-shaped function of the top 1 percent wealth share. Such a long-run relationship resists changes in controls and estimation methods. By considering that in France the top 1 percent currently controls 25% of the total wealth, while in the US, their share is close to 40%, our exercise provides an important insight to policymakers. Redistributive policies targeting the French elite might be desirable for various reasons besides growth. Notably, an increase or a reduction of the share of wealth controlled by the top 1 percent may actually lead to a reduction in growth rates. Conversely, there seems to be plenty of space for wealth redistribution in the US. Bringing down the top 1 percent wealth share from its current levels to approximately 25% can potentially increase the growth rate by 1–2%, which is quite significant considering historical trends reported in Fig. 1.

6. Final considerations

An interest in the relationship between inequality and long-run growth continues to fascinate economists among other social scientists. In recent years, there has been an increasing realisation that wealth is more concentrated than income and that most of the action happens at the very top of the distribution. This study fills a gap in the literature by assessing the correspondence between the wealth share as represented by the top 1 percent and the output growth rate.

The development of the *World Inequality Database* has shown that top income – wealth measures are not just *proxies* for inequality. They are an interesting policy issue in their own right with potentially important political and economic implications. On the one hand, in France, the top 1 percent has an average net personal wealth of approximately US$ 6 million that increases to US$ 70 million as we move to the top 0.01 percent. On the other hand, the magnitudes are even more impressive in the US, with values of approximately US$ 12 million for the first group and approximately US$ 350 million at the very top 0.01 percent. While we acknowledge the intrinsic heterogeneity and mobility of those belonging to these groups, our understanding is that they wield a significant amount of skills and power in society.

Utilising time series cointegration techniques, we study the experience of France and the US between 1950 and 2014. Our estimates are robust to controlling for physical and human capital, degree of trade openness, government size, and asset price movements, using different estimation methods. We show that the growth rate exhibits an inverted-U-shaped curve, as a function of the top 1 percent wealth share. Such a relationship still exists when we narrow the analysis to the top 0.01 percent, though it is no longer statistically significant for France. We also have computed a sort of optimal wealth share, i.e., the share of wealth compatible with a maximum

[7] Under the null hypothesis of cointegration, Hansen's cointegration test proposes the use of the Lc statistic which arises from the theory of Lagrange multipliers for parameter instability. Contrasting with residual based cointegration tests, it does rely on estimates from the original equation in I(1) processes.

growth rate, showing that France is growing close to its long-run potential, while the US is significantly below its. Notably, by reducing wealth concentration to French levels, the US could potentially increase its long-run growth rate by 1–2%.

The main underlying transmission channel lies in the prevailing structure of incentives. The "Rich" form a body of decision makers, so to speak, both in public and private spheres. If they are not adequately rewarded, we might end up with an inefficient allocation of resources. However, extremely high wealth concentration levels are also problematic. In this case, inequality may obstruct the supply of people and ideas into the economy, thereby limiting opportunities for those who are not already at the top. One may also have a subversion of the institutions leading to an ineffective political system and dysfunctional markets.

Institutions are considerably country specific. Our choice of working on the French and US economies allowed us to compare two different institutional arrangements while taking advantage of the time dimension of data. However, a remarkably more equal welfare regime is missing in our exercise, i.e., the Nordic countries. This group is internationally recognised for having lower levels of income and wealth inequality. Future research covering those experiences and perhaps comparing local *optimal wealth shares* is to be encouraged once data become available.

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Appendix A. Empirical appendix

Tables A1 and A2 present a summary of the ADF with and without structural break, and PP unit root tests for France. In Tables A3 and A4, we present the results of the same tests for the US. Outcomes indicate that the output growth rate is stationary while the other series are integrated of order one.

As suggested by one of the reviewers, we also assess the robustness of the inverted-U to the use of NNI rather than GDP. The former is commonly employed in the inequality literature as a better measure for economic development. The NNI series from 1970 to 2014 are obtained from the World Development Indicators. Given that, in this case, we only have data for 44 years – instead of 64 in our previous regressions – we reduce the number of controls to maintain a parsimonious modelling structure. The results are reported in Table A5 and are consistent with our previous findings. The estimated local *optimal wealth share* is slightly lower, ranging between 23% and 24%.

Table A1
Unit root tests (levels) – France.

	y		g	
	Intercept	Trend&Intercept	Intercept	Trend&Intercept
Method	Prob.	Prob.	Prob.	Prob.
ADF	0.1028	0.7882	0.0019	0.0013
PP	0.0268	0.8549	0.0021	0.0013
ADF – structural break	0.0692 (1953)	0.9109 (1959)	<0.01 (1973)	<0.01 (1973)
	ω		ω^2	
	Intercept	Trend&Intercept	Intercept	Trend&Intercept
Method	Prob.	Prob.	Prob.	Prob.
ADF	0.4315	0.8886	0.4085	0.8882
PP	0.3857	0.8643	0.3464	0.8482
ADF – structural break	0.6875 (1964)	0.4508 (1967)	0.4393 (1964)	0.3138 (1967)

Automatic lag selection based on SIC. Newey-West automatic Bandwidth selection.

Table A2
Unit root tests (1st differences) – France.

	y		g	
Method	Intercept Prob.	Trend&Intercept Prob.	Intercept Prob.	Trend&Intercept Prob.
ADF	0.0019	0.0013	0.0000	0.0000
PP	0.0021	0.0013	0.0000	0.0000
ADF – structural break	<0.01 (1973)	<0.01 (1973)	<0.01 (1981)	<0.01 (1981)
	ω		ω^2	
Method	Intercept Prob.	Trend&Intercept Prob.	Intercept Prob.	Trend&Intercept Prob.
ADF	0.0001	0.0002	0.0000	0.0000
PP	0.0000	0.0001	0.0000	0.0000
ADF – structural break	<0.01 (1970)	<0.01 (1996)	<0.01 (1968)	<0.01 (1968)

Automatic lag selection based on SIC. Newey-West automatic Bandwidth selection.

Table A3
Unit root tests (levels) – United States.

	y		g	
Method	Intercept Prob.	Trend&Intercept Prob.	Intercept Prob.	Trend&Intercept Prob.
ADF	0.2211	0.8646	0.0000	0.0000
PP	0.1812	0.8419	0.0000	0.0000
ADF – structural break	0.7141 (1983)	0.8265 (2007)	<0.01 (2006)	<0.01 (1984)
	ω		ω^2	
Method	Intercept Prob.	Trend&Intercept Prob.	Intercept Prob.	Trend&Intercept Prob.
ADF	0.9620	0.9021	0.9759	0.9214
PP	0.9734	0.9021	0.9883	0.9503
ADF – structural break	0.9621 (1996)	0.3567 (1970)	0.9688 (1996)	0.6861 (1969)

Automatic lag selection based on SIC. Newey-West automatic bandwidth selection.

Table A4
Unit root tests (1st differences) – United States.

	y		g	
Method	Intercept Prob.	Trend&Intercept Prob.	Intercept Prob.	Trend&Intercept Prob.
ADF	0.0000	0.0000	0.0000	0.0000
PP	0.0000	0.0000	0.0001	0.0001
ADF – structural break	<0.01 (2006)	<0.01 (1984)	<0.01 (1976)	<0.01 (1976)
	ω		ω^2	
Method	Intercept Prob.	Trend&Intercept Prob.	Intercept Prob.	Trend&Intercept Prob.
ADF	0.0000	0.0000	0.0000	0.0000
PP	0.0000	0.0000	0.0000	0.0000
ADF – structural break	<0.01 (1988)	<0.01 (1971)	<0.01 (2003)	<0.01 (2003)

Automatic lag selection based on SIC. Newey-West automatic bandwidth selection.

Table A5
DOLS estimations using Net National Income.

Dependent variable: NNI_t						
	France			United States		
	XXII	XXIII	XXIV	XXV	XXVI	XXVII
ω_{t-1}	0.072000**	0.142865***	0.159780***	0.733010***	0.314911*	0.522016***
ω^2_{t-1}	− 0.001496*	− 0.003050***	− 0.003378***	− 0.013190***	− 0.006381**	− 0.010856***
α_0	− 0.787170**	18.09478***	− 0.743358	− 10.80407***	− 237.7701***	− 246.9223***
α_1	0.017499***	0.027286***	0.017151**	0.051113	− 0.341102	− 0.384853***
NNI_{t-1}	0.594601***	0.613471***	0.364357***	0.921581***	0.780916***	0.869096***
$Pop._{t-1}$	–	− 5.580711***	− 0.722973	–	49.34866***	56.26388***
Capital $stock_{t-1}$	–	− 0.045097	− 0.014000	–	− 2.539587	− 8.126697***
HC_{t-1}	–	6.732899***	7.016693***	–	− 1.135950	1.343146
Gov_{t-1}	–	–	− 2.191985***	–	–	17.63808**
$Openess_{t-1}$	–	–	− 0.005497***	–	–	0.037854*
Time dummies	✓	✓	✓	✓	✓	✓
Lc stat.	0.143998	0.367638	0.391435	0.087490	0.205007	0.208963
Jarque Bera	0.832347	1.213420	3.809217	0.144961	0.404611	0.281307
ω^*	24.06	23.42	23.65	27.78	24.67	24.04

*, **, *** stand as 10%, 5%, and 1% of significance.

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