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JEL Classification: C23, D31, E21, O5

Keywords: Saving, income distribution, panel data, non-linearities

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Income Distribution and Aggregate Saving: A Non-Monotonic Relationship

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August 2016

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1 Introduction

Is there an empirical link between income distribution and aggregate saving? Drawing on the experience of 29 OECD countries, this paper suggests *yes*, but in a non-monotonic way. It suggests that at a low level of inequality, more inequality is associated with higher saving, but it also shows that a negative relationship between inequality and saving prevails at high levels of inequality.

Given the secular rise in income inequality, economists increasingly focus on the macroeconomic implications of this development. A link between inequality and saving lies at the heart of this literature: For instance, the debate about secular stagnation has drawn new attention to the Keynesian idea that rising inequality increases the aggregate propensity to save and thus exerts a drag on aggregate demand (e.g., Eggertsson and Mehrotra, 2014; Summers, 2015). Assuming the same positive relationship between inequality and saving, but coming to a different conclusion, the neoclassical growth literature suggests that inequality promotes economic performance by fostering capital accumulation (Bourguignon, 1981). Yet, with regard to global current account imbalances, some studies argue that an increase in inequality lowers private saving and the current account (Ranciere et al., 2012; Al-Hussami and Remesal, 2012; Behringer and van Treeck, 2013).

Although household saving constitutes a common transmission variable in all these strands of literature, the link between income inequality and saving is theoretically and empirically unclear: As richer households tend to have a higher propensity to save than households at the lower end of the income distribution (e.g., Dynan et al., 2004), an increase in income inequality may cause a rise in aggregate saving (Keynes, 1936, 1939). Yet, if households engage in upward-looking interpersonal comparison, middle- and low income earners might lower their saving rate in response to rising top incomes (Drechsel-Grau and Schmid, 2014; Bertrand and Morse, 2016). Thus an increase in inequality could just as well trigger expenditure cascades and a decline in aggregate saving (Alvarez-Cuadrado and El-Attar Vilalta, 2012; Frank et al., 2014).

In line with the theoretical ambiguity, cross-country and panel-data studies that investigate the effect of inequality on national or private saving rates often remain inconclusive (Schmidt-Hebbel and Serven, 2000; Li and Zou, 2004; Leigh and Posso, 2009). With regard to household saving some studies find a negative effect of inequality, albeit they rely on samples of rather few countries (Leigh and Posso, 2009; Alvarez-Cuadrado and El-Attar Vilalta, 2012; Behringer and van Treeck, 2013).

The present study is the first that focuses on the saving rate of the household sector, while profiting from a large and consistent dataset. By combining saving rates from OECD databases with net income Gini coefficients from the Standardized World Income Inequality Database (SWIID), we construct an unbalanced panel of up to 792 observations from 29 advanced economies.[1] Following Atkinson and Brandolini (2001, 2009), the SWIID maximizes data comparability in terms of common income measures and household units (see Solt, 2015). Yet, given the remaining data uncertainty, we test for the robustness of our results by running multiple imputation regressions; and by applying alternative inequality data from the Luxembourg Income Study and the World Top Incomes Database.

In consistence with the theoretical ambiguity and the inconclusiveness of the empirical literature, we do not find a linear correlation between inequality and saving. However, we extend the literature by exploring non-linearities. We reveal a highly significant hump-shaped relation between inequality and saving that is robust to a large set of controls, including equity and house prices, credit availability, and financial liberalization. We find that the impact of inequality on saving is positive at low levels of inequality, whereas it becomes negative after some turning point, roughly located at a Gini between 28 and 30. This hump-shaped pattern is robust to endogeneity concerns, different estimation techniques, measures of inequality, and sample compositions. Yet the pattern appears to have vanished in the aftermath of the financial crisis.

As the availability of credit financing might be a precondition for expenditure cascades (see, e.g. Rajan, 2010, Frank et al., 2014, Bertrand and Morse, 2016), we also test whether the impact of inequality interacts with credit availability and financial market liberalization. We find that rising inequality tends to reduce saving if financial markets are widely liberalized or the ratio of credit to GDP is high. Nonetheless, in both a low-credit and high-credit environment, the hump-shaped relationship between inequality and saving prevails.

While we primarily focus on household saving rates, we find some evidence that the hump-shaped effect of inequality also appears for private saving rates, national saving rates, and the current account balance.

The paper proceeds as follows: Section 2 describes the theoretical background to the analysis. Section 3 briefly reviews the recent empirical literature on the household, state, and cross-country level. Section 4 describes the data, focusing on measures of saving and income distribution. Section 5 reports our baseline regression results, followed by an extensive sensitivity analysis, an exploration of interaction effects, and regressions for alternative dependent variables. Section 6 discusses the results and concludes.

[1] We use version 5.0 of the SWIID (Solt, 2016), which is a major update of the original database summarized in Solt (2009).

2 Theoretical link between income distribution and household saving

The link between income distribution and aggregate household saving is ambiguous, as there are various opposing effects on the microeconomic level, which might be offsetting in the macroeconomic aggregate: First of all, according to Keynes (1939), the individual propensity to consume decreases with personal income, which implies "[...] that the collective propensity for a community as a whole may depend (inter alia) on the distribution of incomes within it." Possible explanations for higher saving rates of richer households are bequests or wealth that enter the utility function as luxury goods (e.g., Carroll, 1998). Moreover, asset-based means testing for social security benefits (e.g., Hubbard et al., 1995; Gruber and Yelowitz, 1999) and a subsistence consumption level lying above the income of poorer households (Musgrove, 1980) can lower the saving rates of poorer households.

Following these arguments for a positive relationship between individual incomes and saving rates, a rising concentration of income at the top should lead to a rise in the aggregate saving rate. However, if consumption or saving decisions of different households are mutually interrelated, the opposite can be true: According to the relative income hypothesis "[...], the frequency and strength of impulses to increase expenditure for one individual depend entirely on the ratio of his expenditures to the expenditures of those with whom he associates." (Duesenberry, 1949, p. 32).

Building upon such consumption externalities, Frank et al. (2014) propose a formal model of "expenditure cascades". Similarly, Alvarez-Cuadrado and El-Attar Vilalta (2012) incorporate relative income considerations into an OLG model. In both models, increasing consumption of a reference group encourages additional consumption by households further down the income ranking. On aggregate, a mean preserving spread in incomes thus leads to a decrease in the saving rate.[2]

In conclusion, the prerequisite for a decline in the aggregate saving rate due to rising inequality is that saving rates of low and middle income earners decline sufficiently; so that the increase in the volume of saving, resulting from the shift in income toward households with a larger propensity to save, is overcompensated. To enable this decline in saving, the

[2] A decline in the aggregate saving rate can also result from a decline or stagnation of income at the bottom of the distribution. According to the habit persistence theory (Brown, 1952), people lower their saving rate to hold on to their usual consumption level when real income deteriorates. If people are used to steady improvements in living standards, habit persistence may thus implicate lower saving when income growth slows down for certain income groups. Similarly, a decrease in aggregate saving can result when more and more households are falling below a subsistence consumption level. The latter would be most pronounced, if the subsistence level is a socially acceptable consumption standard that is high enough to affect a large number of households.

initial saving rates (or the financial wealth) of low and middle income households have to be sufficiently large. Otherwise, if saving rates (and wealth) are already low, poorer households have to borrow to finance their excess consumption.

3 A brief survey of the empirical literature

The link between income distribution and household saving has been tested in a couple of micro- and macro-data studies. Using survey data from the U.S., a highly cited study by Dynan et al. (2004) finds a strong positive correlation between saving rates and household incomes. Yet, based on Canadian data, Alan et al. (2015) indicate that saving rates do not differ substantially across long-run income groups. Like Dynan et al. (2004), Alvarez-Cuadrado and El-Attar Vilalta (2012) find that saving rates increase in permanent income. Moreover, the latter study emphasizes a negative correlation between the income growth of local reference groups (or an increase in inequality) and the saving rates of poorer households. Similarly, Bertrand and Morse (2016) support the relative income hypothesis and "trickle-down consumption" by showing that middle income households consume a larger share of their income when exposed to higher upper income and consumption levels. Based on this result they estimate that in 2005 the aggregate personal saving rate in the US might have been 1.1 to 1.3 percent higher, if income growth at the top had not outpaced growth at median levels. Finally, Drechsel-Grau and Schmid (2014) show that "keeping up with the Joneses behaviour" is not limited to one side of the Atlantic. Using data from the German Socio-Economic Panel they find that an increase in reference consumption by 1% leads households to raise their own consumption by about 0.3%.

Altogether, micro-data evidence supports both the Keynesian- and the relative income hypothesis. Yet it says little about aggregate saving because it cannot tell which of the opposing effects prevails. Therefore we have to refer to macro-data studies, which regress aggregate saving rates on aggregate measures of income distribution.

In general, cross-country studies on inequality and saving often remain inconclusive and the results vary with the estimation technique and sample composition. Because of data restrictions either national- or private saving rates serve as the (main) dependent variable in most macro-data studies. To provide a better comparability within the literature and to our own paper, we restrain our survey to panel regressions and subsamples of data from developed economies or OECD members. Drawing on this selection, Schmidt-Hebbel and Serven (2000), Li and Zou (2004), as well as Leigh and Posso (2009) do not find a consistent relationship between inequality and saving. Smith (2001), however, reports a positive effect of inequality on private saving.

To our knowledge, there are only three studies that (also) examine the effect of income distribution on the saving rate of the household sector. Regressing household saving on lagged top income shares, in a panel of 10 developed economies observed between 1975 and 2002, Leigh and Posso (2009) find no significant effect of inequality. In contrast, Alvarez-Cuadrado and El-Attar Vilalta (2012) suggest a negative impact of inequality on aggregate saving. Drawing on a sample of 6 developed economies, observed between 1954 and 2007, they find a negative effect of the top 5% income share, which is highly significant under a range of different econometric specifications. A recent study by Behringer and van Treeck (2013) primarily deals with the effect of income distribution on the current account. Yet it also takes a look at saving rates and financial balances of the household sector. In a sample of G7 economies, the study finds a significant negative effect of the top 5% income share, while the Gini coefficient appears to be insignificant.

Altogether, the literature about the relationship between inequality and saving remains inconclusive, which might be due to some deficiencies: First, there are only few studies that examine the aggregate saving rate of the household sector. Second, the studies which focus on household saving are based on very few countries. Third, the existing literature does not control for a number of covariates, which could lead to an omitted variable bias; and fourth, it does not account for a non-monotonic relationship.

4 Data description

4.1 Saving rates and sample composition

Most existing studies focus on national saving, which measures the total amount of saving in the economy, including households, firms and the government. Yet, since most theories about saving and inequality refer to household behavior, we prefer to focus on household saving rates, while we will glance at broader measures of saving and the current account balance at the end of this paper.

Although household saving rates are less readily available than national saving rates, we are able to compose a fairly large sample by combining data from the OECD National Accounts Database with data from the OECD Economic Outlook. Thus we receive an unbalanced panel of up to 792 annual observations from 29 countries. Our panel is highly unbalanced, with the earliest observation being from 1961 and the latest from 2013. To benefit from a homogenous sample of high quality data, we limit our panel to high-income OECD countries, as defined by the World Bank classification.

The OECD calculates saving by subtracting household consumption expenditures from household disposable income, net of fixed-capital depreciation. Capital holding gains are not included, which is conducive to our focus on active saving behavior. Dividing the saving volume by the disposable income of the household sector yields the saving rate.[3]

4.2 Inequality data

Our primary measure of inequality is the net Gini coefficient (post-tax, post-transfer), sourced from version 5.0 of the Standardized World Income Inequality Database (SWIID) generated by Frederic Solt (2009, 2016). Following the recommendations from Atkinson and Brandolini (2001, 2009) the SWIID aims to provide the most comparable data for the broadest possible sample of countries and years (see, Solt, 2015). To achieve this goal, it collects Ginis from a large number of sources like cross-national inequality databases, national statistical offices, and scholarly articles. Then, it adds market and net Ginis from the Luxembourg Income Study (LIS) as a benchmark of most reliable data. As the Ginis from the source data are often not directly comparable due to different income definitions or accounting units, the SWIID uses a multiple-imputation algorithm to estimate standardized net and market Ginis for all country-years not yet covered in the LIS. To reflect the uncertainty associated with these estimates, the SWIID 5.0 reports 100 imputations for each observation, generated via Monte Carlo simulations.

There are two alternative paths to employ the SWIID data in regression analysis. The first is to average the imputations and to use the resulting point estimates with usual regression techniques, thereby simply ignoring the uncertainty in the inequality data. The second, which is recommended by the author of the SWIID, is to deploy multiple imputation tools that explicitly account for data uncertainty within the estimation results.

As the uncertainty related to Ginis from high-income OECD countries is relatively low, this paper primarily uses point estimates of the SWIID data. However, we also employ multiple imputation estimation techniques, and inequality data from alternative sources, to test for the robustness of our results.

4.3 Control variables

To isolate the true impact of income distribution on saving, we control for a number of variables that are so far neglected in the literature on inequality and saving. First of all, we are concerned about wealth effects being a cause of spurious regressions. Rising asset prices may cause a drop in saving, as people feel wealthier and are able borrow against higher collateral

[3]Notably, the household sector includes unincorporated enterprises and in most cases also non-profit institutions serving households.

(e.g. Slacalek, 2009; Hüfner and Koske, 2010). However, if an asset bubble is associated with growing income inequality, these wealth effects may misleadingly be attributed to income distribution. To avoid such an omitted variable bias, we employ an indicator of real house price developments (*houses*) and real stock market returns (*equities*).

Another potentially important control is the availability of credit, which we proxy with the ratio of private credit to GDP (*credit*). Whereas financial liberalization may enhance saving opportunities, a greater availability of credit could as well boost private consumption by relaxing borrowing constraints (e.g., Bandiera et al., 2000). As an expanding financial sector may affect income distribution (e.g., Delis et al., 2014; Bumann and Lensink, 2016), omitting financial depth may cause a bias in the estimated effect of inequality.

The remaining control variables are common in the literature on inequality and saving. The old-age dependency ratio (*depend*) is defined as the share of population aged 65 or older over the working-age population. According to the life-cycle hypothesis (Modigliani, 1970) we expect a negative sign for its estimated coefficient. The variable *incgrow* denotes the growth rate of households' real disposable income per capita.[4] Because of habit persistence an increase in income may lead to an increase in saving. However, if households are forward looking, consumption may also rise in anticipation of rising future incomes. Real interest rates are measured by the real return on long term government bonds (*interest*). Although in standard macroeconomic models a higher interest rate increases the attractiveness of saving compared to consumption, the sign of its effect is ambiguous. If households pursue a fix amount of savings, higher interest rates could as well reduce saving because less money must be put aside to reach a saving target. Further controls are the fiscal balance (*fiscal*), to account for Ricardian equivalence; the natural logarithm of GDP per capita (*ln(gdppc)*); and the inflation rate (*infl*). A more detailed description of the sources and derivations of our variables can be found in the Appendix. Table 1 contains summary statistics.

Finally, the saving rate of private households is likely to be affected by factors that are unobservable or difficult to measure. For instance, cultural attitudes (like the proneness for competitive thinking) could be a source of omitted variable bias, if they affect attitudes towards consumption as well as the political stance towards redistribution.[5] To control for such time invariant factors our baseline model includes country fixed effects.

[4] We prefer the growth rate of household disposable income over the GDP growth rate, due to less severe concerns about reverse causality and its more direct impact on the household sector.

[5] Catte and Boissinot (2005) emphasize further factors, which could explain differences in household saving rates. These include the number of unincorporated enterprises in the household sector, the provision of public goods, the role of direct versus indirect taxation, and the design of the pension system. However, after adjusting the data for differences in public provision and the tax system, Catte and Boissinot (2005) find only modest effects on the level and international differences in saving rates.

[Table 1 about here]

5 Empirical findings

We now turn to the empirical assessment of the relationship between inequality and household saving. First, we present our regression model along with our baseline results. Next, we show that the results are robust to endogeneity, data uncertainty, alternative inequality measures, different sample compositions, and a flexible functional form. Then, we test whether the relationship between inequality and saving interacts with financial market conditions. Finally, we analyze the effect of inequality on some broader measures of saving as well as the current account balance.

5.1 Baseline results: A hump-shaped relationship

Table 2 shows the results of our baseline regressions. Based on an unbalanced panel of up to 792 observations from 29 advanced economies we estimate the following equation:

$$\text{saving}_{it} = \alpha + \beta_1 \text{gini}_{it} + \beta_2 \text{gini}_{it}^2 + \beta' \text{X}_{it} + \alpha_i + \lambda_t + \epsilon_{it} \tag{1}$$

where $saving_{it}$ is the aggregate saving rate of the household sector in country i and year t.[6] Among the regressors we focus on the Gini of net incomes, which we include in a linear and a squared form, to allow for a non-linear relationship. The vector X_{it} denotes our set of control variables; α_i are country fixed effects; λ_t are time fixed effects; and ϵ_{it} stands for the error terms.

In Table 2 each pair of columns reports two identical models, which only differ by the inclusion of the quadratic term of the Gini in the even numbered columns. Columns (1) and (2) report pooled OLS estimates, whereas Columns (3)-(8) contain country fixed effects. Above all, Table 2 shows that adding a quadratic term of the Gini coefficient to a model of household saving reveals a concave relationship between inequality and saving. Whereas the pooled OLS models indicate no significant correlation between inequality and saving, in all fixed effects models the Gini and its square are highly significant, when included together.

In Columns (3) and (4) we exploit the maximum number of available observations by focusing on small fixed effects models, which include only the most essential control variables. While the linear specification in Column (3) shows no systematic effect of inequality, the

[6] To deal with gaps in the data and to weaken serial correlation in the residual, some previous studies consolidate the annual data into 5-year averages. Our regression results are very similar with averaged data (available upon request). Yet we prefer the use of annual data as most of the benefits of averaging are obsolete with our dataset and the use of cluster robust standard errors.

estimated coefficients of *gini* and $gini^2$ in Column (4) indicate a hump-shaped function between saving and inequality.

To asses the statistical significance of this nonlinear relationship, we report the results of the Sasabuchi-Lind-Mehlum (SLM-Test) in the lower part of Table 2 together with the slopes at the minimum and maximum values of the Gini in our sample.[7] In addition, we report the Fieller 90% confidence intervals for the turning points. With a p-value of 0.005 the SLM-Test rejects the null of a monotone or U-shaped relationship in favor of an inverted-U-shaped (concave and hump-shaped) relationship. The turning point, at which the marginal effect of inequality becomes negative, is estimated at a Gini of 28 within a 90% confidence interval of 25 to 33. Thus the turning point corresponds roughly to the median value of the Gini in our regression sample.

[Table 2 about here]

[Figure 1 about here]

Based on the results of Column (4), Figure 1 illustrates the marginal effect of inequality on saving across different values of the Gini. It pictures how the effect of inequality is decreasing with an increasing level of inequality. The marginal effect of inequality ranges from 0.85 at the smallest Gini in the sample (Gini of 18, observed in Sweden 1990) toward -1.79 at the upper bound (Gini of 49, in Chile 2009). In line with the results from the SLM-test, the confidence intervals reveal a significantly positive effect of inequality for Ginis ranging from 18 to 25 and a significantly negative effect for Ginis above 33.

In the previous section we suppose that wealth effects and credit conditions may cause a downward bias of the estimated effect of inequality on saving. To account for this bias, Columns (5) and (6) report extended models which control for two measures of asset price movements (*equities* and *houses*) and the credit to GDP ratio (*credit*). Following preceding studies, we additionally include the log of real income per capita (*ln(gdppc)*) and the inflation rate (*infl*).[8]

Altogether, the inclusion of the new controls hardly changes the relationship between inequality and saving, although increasing asset prices indeed correspond to slightly lower

[7] The SLM-Test was developed by Lind and Mehlum (2010), based on the work of Sasabuchi (1980).

[8] The introduction of the new control variables slightly alters the composition of the regression sample, primarily due to the loss of observations from earlier time periods. Moreover, the range of inequality levels narrows, as Chile, which is marked by extreme levels of inequality, is not included in the remaining sample of 616 observations. Thus the maximum value of the Gini in the sample decreases to 37.8, while the minimum value is still 18.

saving rates. While the average effect of inequality turns positive in the linear model of Column (5), it is still insignificant. In the quadratic model of Column (6), however, the Gini and the square of the Gini are again highly significant. With the new regressors and the altered sample, the slope at the lower bound increases to 1.3 and the slope at the upper bound decreases to -2, indicating that the hump-shaped relationship becomes somewhat steeper.[9] Moreover, the estimated turning point shifts to the right and is now positioned at a Gini of 30, within a 90% confidence interval of 27 to 34.[10]

Finally, in Columns (7) and (8) we add year-dummies in order to account for common shocks, like the global financial crisis. Although few of the dummies are significant, their introduction slightly affects the estimates of other variables, like equity prices, which become insignificant. Nonetheless, for $gini$ and $gini^2$ the results remain almost unchanged, yielding a hump-shaped relationship with a turning point at a Gini of roughly 29 and a 90% confidence interval ranging between 27 and 32.[11]

In sum, our regression models resemble earlier studies, which do not find a linear relationship between income inequality and the aggregate saving rate. However, by introducing a quadratic term, we reveal a hump-shaped relationship, which peaks at a net Gini roughly between 28 and 30. To get an idea of the countries that have driven the non-linear effect, Figure 2 plots the Ginis observed in 1995 along with the associated marginal effects.[12] Looking at the two polar cases, the figure predicts a strongly positive effect of rising inequality on saving in Sweden and a negative effect in the United States. In the following section we test for the robustness of the hump-shaped relationship.

[Figure 2 about here]

[9] To allow for comparability between different models, we report the slopes at the boundaries of the full sample, i.e. at Ginis of 18 and 49.

[10] When we run regression (4) on the reduced sample of 616 observations, the estimated coefficients of $gini$ (3.458) and $gini^2$ (-.058) in the small model are more similar to the ones from Column (6). The turning point is then estimated at 29.6.

[11] Following Grigoli et al. (2014) and Loayza et al. (2000), we also add the share of urban population, terms of trade, and the young age dependency ratio as additional regressors. Whereas the latter two variables are positively related to saving, the results for $gini$ and $gini^2$ are almost unchanged by this exercise. Finally, the concave relationship is also robust to the fixed effects model from Schmidt-Hebbel and Serven (2000), who control for young- and old-age dependency, gdp growth, per capita GDP and also the square of per capita GDP. Results are available upon request.

[12] Corresponding figures for different time periods are available upon request. We have chosen to picture the marginal effects in 1995 as it constitutes a time period that stands rather at the beginning of the sample, but already contains most of the countries.

5.2 Robustness Tests

5.2.1 Addressing endogeneity via lag identification, 2-SLS and System GMM

Drawing on economic theory, we have so far assumed a causal relationship running from the explanatory variables towards saving. Yet in particular the estimated coefficients of disposable income growth, the interest rate, and the fiscal deficit could be subject to simultaneity or reverse causality. But also the estimates of the dependency ratio and inequality, which are less likely to be directly affected by household saving, may be biased from the endogeneity of other variables. This section addresses this problem from several angles.

In Column (1) of Table 3 we follow the simplest approach for causal inferences in a panel setting by using lagged instead of contemporaneous values of the explanatory variables. The results are almost identical to the results from estimations with contemporaneous regressors, confirming the hump-shaped relationship with a peak value that is roughly located at a Gini of 28. The same is true when we vary the lag length between 2 and 5 years (results are available upon request), similarly to the approach taken by Leigh and Posso (2009).

[Table 3 about here]

Our second approach to account for possibly endogenous regressors is a 2-SLS (two-stage least squares) estimation, reported in Column (2), where *interest* and *fiscal* are instrumented with their first and second lags; and *incgrow* is instrumented with the first and second lag of real gdp growth.[13] Whereas the coefficient of incgrow becomes considerably smaller and insignificant with this approach, the results for all other variables are hardly affected. Above all, *gini* and $gini^2$ remain highly significant, still indicating a hump-shaped relationship with a peak at a Gini of 28.6.

While we first simply assumed that the Gini is exogenous, we also test for the validity of this assumption via a difference-in-Sargan statistic. By not rejecting its null hypothesis (p-value: 0.76), the difference-in-Sargan statistic confirms the exogeneity of *gini* and $gini^2$. Moreover, a test for the null that the presumably endogenous regressors (*incgrow*, *fiscal* and *interest*) can actually be treated as exogenous yields a p-value of 0.163, indicating that it is reasonable to rely on simple fixed effects models for the baseline regressions.

Another way to control for reverse causality is the construction of a dynamic model with a lagged dependent variable that partially captures feedback effects, running from past saving

[13] We deploy the xtivreg2 stata routine by Baum et al. (2003) and Schaffer (2010) to estimate a 2-SLS model with cluster-robust standard errors. The test statistics tell us that the instruments are both relevant and orthogonal to the error term. Above all, the Hansen J test does not reject its null of the instruments orthogonality (p-value: 0.89), whereas the Kleibergen-Paap rk LM statistic rejects the null of underidentification (p-value: 0.03). The instruments' relevance is underlined by the Kleibergen-Paap rk Wald F statistic of 7.32, which suggests a maximal relative IV bias of roughly 10%.

towards current values of the explanatory variables. Yet, as demonstrated by Nickell (1981), in fixed effects models the coefficient of the lagged dependent variable y_{it-1} is correlated with the error term and thus introduces a downward bias, which decreases with the length of the time dimension.

One solution for this problem is the adoption of a system GMM model, similar to the one deployed in Loayza et al. (2000) and Grigoli et al. (2014). The system GMM estimator by Blundell and Bond (1998) and Arellano and Bover (1995) is an advancement of the difference GMM estimator (Arellano and Bond, 1991), which differences the saving equation to discard the fixed effects and then uses second and higher lags of the dependent variable as instruments for $y_{it-1} - y_{it-2}$. To account for the possible endogeneity of additional regressors, difference GMM adopts second and higher lags of X_{it} as instruments for $X_{it} - X_{it-1}$. One weakness of difference GMM is its poor performance in finite samples and with persistent dependent variables. To solve this problem the system GMM estimator adds an additional equation in levels, thus building a system of two simultaneous equations. For the levels equation lagged first differences are used as instruments, assuming that the additional instruments are orthogonal to the fixed effects.[14]

Because of its greater reliability in smaller samples, we use the one-step version of system GMM with cluster robust standard errors. To avoid an over-fitting of endogenous variables with too many instruments, resulting from the relatively long time dimension, we use a collapsed instrument matrix (see, Roodman, 2009) and restrict the instruments for the transformed equation to lag 2 and lag 3. To maximize the sample size in our unbalanced panel, we use orthogonal deviations instead of the first difference transformation. In line with Loayza et al. (2000) we treat *incgrow*, *interest*, *fiscal*, *credit*, *infl* and *ln(gdppc)* as endogenous, while we consider the dependency ratio as exogenous. Moreover, we treat *gini*, *gini*^{2}, *equities* and *houses* as endogenous variables.

Columns (3)-(5) of Table 3 present the results from our system GMM estimations.[15] Whereas the large and highly significant point estimate for the lagged saving rate in each model indicates a high degree of persistence, the estimated coefficients of the other regressors are considerably smaller than in the static models. However, in dynamic models the coefficients of the saving determinants only capture short-run effects, which are difficult to measure because of the large share of variation captured by the lagged dependent variable.

[14] Although the system GMM estimator was actually designed for panels with a large number of cross-units, Soto (2009) demonstrates that the estimator also works reasonably well when N is small.

[15] Standard specification tests for system GMM are given at the bottom of the table. Most importantly, the AR(2) p-values confirm the model specification by never rejecting the null of no second order auto-correlation in the error term. However, the Hansen-J-test rejects its null at the 10% level in Column (3), which may cast doubt on the validity of the instruments. In columns (4) and (5) the Hansen test is far from rejecting the null, but it is not fully reliable due to the large number of instruments (see, Roodman, 2009).

The point estimates of lagged saving (0.9 or larger) indicate that the long-run effect of a permanent change in the saving determinants is roughly 10 times the size of the short run effect. For the short-run effects of *gini* and $gini^2$ a significantly positive respective negative sign appears in the small model of Column (3), but vanishes after the introduction of the year dummies (Column 4) and the additional regressors (Column 5). Yet the SLM-tests indicate that a hump-shaped relationship between inequality and saving is significant, with a turning point at a Gini of roughly 29, in Columns (3) and (4); and just marginally below significance, with a somewhat lower peak value, in Column (5).

Altogether, the results from lag identification and 2-SLS models fully confirm the results from standard fixed effects models, whereas system GMM indicates effects of inequality that are fairly small in the short run and associated with a high degree of uncertainty due to the strong persistence of the saving rate. Since our interest lies foremost in the medium to long-run relation between inequality and saving, we will continue with static models.

5.2.2 Multiple imputation estimations

In this section we test whether the uncertainty that is associated with the SWIID data affects our results. Therefore, we follow the advise from Solt (2016) and employ a multiple imputation technique to account for data uncertainty. Essentially, Stata's multiple imputation estimation routine, which we apply in this section, runs repeated regressions for each of the 100 imputations of the net Gini and then pools the resulting estimates following the combination rules proposed by Rubin (1987). Thus the estimated coefficients and standard errors are adjusted for the variability between imputations, whereas regressions on averaged data treat the Gini from the SWIID as an error-free variable.[16]

[Table 4 about here]

Table 4 shows that using multiple imputations instead of averaged data only marginally affects our results. To provide direct comparability, each regression exactly resembles the quadratic models of the baseline specification, but is estimated with the multiple imputation technique. Just like in the baseline table we find a hump-shaped relationship between inequality and saving. The effect of inequality remains highly significant and the locations of the turning points almost unchanged, although the estimated coefficients become somewhat smaller and the standard errors slightly larger with multiple imputations.[17] Altogether,

[16] Brownstone and Valletta (2001) offer an excellent summary of the multiple estimation technique and its applications in economics.

[17] The resulting slightly decreased standard errors together with flattened regression lines are surprising, given that we would normally expect that multiple imputation estimations increase the standard errors. We

the enhanced statistical accuracy stemming from multiple imputation estimates hardly affects our results, which means that we can safely proceed with less computational intensive regression techniques.

5.2.3 Alternative inequality measures

Another approach to account for data uncertainty is the use of alternative inequality measures. The estimates in Columns (1)-(3) of Table 5 are based on inequality data from the Key Figures of the Luxembourg Income Study (LIS). The LIS is often regarded as the gold standard for harmonized inequality data, but its downside is a strongly restricted data availability. By dispensing all information from country-years where uniform income data is not directly available, the regression sample shrinks to a maximal number of 143 observations. While 25 countries remain in the panel, there are only 5 observations available for each country on average. Moreover, mostly observations from countries with extreme levels of inequality are lost when switching from the SWIID to the LIS, so that the boundary values are now located at Ginis of 19.7 and 37.7.

Based on the LIS data, the regression coefficients again indicate a highly significant hump-shaped relationship, both in the the pooled OLS (Column 1) and the fixed effects (Columns 2) model. Compared to the baseline results, the turning points shift slightly to the right, reaching a Gini of 28.8 in in the fixed effects model. Yet, given the substantially different sample composition, the results are similar enough to conclude that the hump-shaped relationship is robust to alternative data sources.[18] Moreover, a hump-shaped relationship also occurs in Column (3), where we deploy the 80/20 percentile ratio from the LIS Key Figures as an alternative indicator of income distribution.

[Table 5 about here]

A general problem with inequality data from income surveys is differential non response (see, e.g. Atkinson and Brandolini, 2001), which is why the development of top incomes is possibly not fully reflected in survey-based inequality measures like Ginis from the LIS or SWIID. To address this problem, we deploy top income shares from the World Top Incomes Database (WTID) by Atkinson et al. (2015) as an alternative inequality variable. Being generated by tax collecting agencies, the WTID data could be more reliable than survey data.

are grateful to Frederic Solt for pointing out a possible explanation: In cases where influential outliers with large standard errors are pulling up the coefficients, using multiple imputations may flatten the coefficients and also estimate them with more precision.

[18] Running the other regression models of Table 2 with the LIS data yields results that are similar to those received with the SWIID data. Results are available upon request.

However, there are also some limitations (see, Atkinson et al., 2011): First, top income shares do not reflect distribution within the middle and lower ranges of the income ranking. Second, the data is based on gross incomes and ignores governmental redistribution. Third, due to diverse tax bases, income definitions, and units of observation the data is not comparable across countries, which is why the data should not be used with estimators that exploit cross-country variations.[19]

Nonetheless, measuring inequality via top income shares could also be worthwhile from a theoretical perspective. The expenditure cascades model proposes that a rapid growth of top incomes provokes a decrease in saving by middle class households, leading to a decline in aggregate saving. Thus an increase in the top income share might exert a more negative effect on the aggregate saving rate than a rise in overall inequality, which is captured by the Gini coefficient.[20]

Column (5) shows the results of our baseline quadratic fixed-effects model with the top 1% income share as the main explanatory variable. With the new measure of inequality the hump-shaped relationship between inequality and saving becomes insignificant. Yet also in the linear model specification of Column (4) the estimated coefficient of the top income share is insignificant. Similarly, when we regress *saving* on $top1inc$ while controlling for *gini* and $gini^2$ in Column (6), the estimated effect of the top income share is not significantly different from zero. For the Gini, however, a hump-shaped relationship persists, which is why we can rule out that the insignificance of the top income share merely results from the altered sample composition.

Altogether, we do not find a significant relationship between top income shares and saving, although our results are not necessarily conflicting with the micro-econometric evidence for trickle-down consumption and expenditure cascades. Compared to the linear regression models of the baseline table, the negative effect of inequality is more pronounced when inequality is measured with top incomes instead of Gini coefficients. Yet the results are insignificant with both measures of inequality, indicating that expenditure cascades are apparently not strong enough to dominate the overall effect.

[19] In the full WTIID database there are also some breaks within countries due to changes in tax legislation etc. When compiling our panel we took care to employ homogenous series for all countries, which leads to shorter time dimensions in Finland and the UK (see, Data Appendix).

[20] Van Treeck (2014) suspects differential effects of Ginis and top income shares on financial stability and personal debt-to-income ratios. Behringer and van Treeck (2013) find differential effects on the current account balance.

5.2.4 Different sample compositions

In this section, we test whether the non-monotonic relationship between inequality and saving is robust to variations in the sample composition, in addition to the reduction in sample size that results from the use of the larger regression model. First, we eliminate the top and bottom 5% of the distribution of saving and inequality from the regression sample. As it can be seen in Column (1) of Table 6, the hump-shaped relationship is robust to the omission of these outliers and also the turning point is still roughly located at a Gini of around 30.

[Table 6 about here]

Next, we strongly limit our sample along the cross-sectional dimension by only including the G7 economies in Column (2). A similar sample has been used in previous studies (see, Alvarez-Cuadrado and El-Attar Vilalta, 2012; Behringer and van Treeck, 2013), which found a negative effect of the top income share. Yet, with our inequality data and the inclusion of the additional control variables, no clear effect of inequality emerges within the G7 sample. Whereas the signs of *gini* and $gini^2$ hint towards a hump-shaped relationship, the effects are far from significant. Possibly this is due to the reduced efficiency, stemming from the very narrow sample.

In further regressions (available upon request) we test for the omission of single countries (one at a time) from the regression sample. In all these regressions a hump-shaped relationship always remains significant and the inequality turning-point varies little.

Moreover, we also check whether the effect of inequality is sensitive to different time periods. When the actual function between inequality and saving is quadratic, the estimated coefficient of inequality is downward biased in a linear regression equation. Yet the bias is small when the regression sample contains only few observations with high values of inequality.[21] As there may have been fewer instances of high inequality, this could explain why Smith (2001) has found a monotonic positive effect within the 1960-1995 period. To test for this supposition, Column (3) reports a regression that only draws on observations from the period 1961-1995.[22] Yet the model yields clear evidence for a hump-shaped relationship, which peaks at a Gini of roughly 28.[23]

In Columns (4)-(6) we continue with restricting the sample along the time dimension. More precisely, we subsequently eliminate the oldest observations, starting with the 1970s

[21] In the context of finance and growth, Arcand et al. (2015) offer a detailed description of the bias in linear models when the true relationship is non-monotonic.

[22] We rely on the small regression model in order to utilize the observations from the 1960s because some of the controls from the large model are not available before 1970.

[23] In a standard linear regression equation the effect of inequality is insignificant (Coef. .010; SE. .169).

in Column (4), the 70s and 80s in (5) and finally also the 90s in (6). In the first two samples the hump-shaped relationship remains highly significant and the turning-point becomes somewhat larger. Only in the sample that solely draws on the most recent observations no significant effect occurs.

5.2.5 Semiparametric regressions

In this section, we allow inequality to take a flexible functional form by estimating a semiparametric regression model:

$$\text{saving}_{it} = f(gini_{it}) + \beta' \mathrm{X}_{it} + \alpha_i + \epsilon_{it} \tag{2}$$

where the control variables X_{it} enter the model linearly and $f(gini_{it})$ denotes an unknown function of the Gini.

Our panel data regressions are based on Baltagi and Li (2002), whose estimator was build into *Stata* by Libois and Verardi (2013). Essentially, the estimator relies on a first difference transformation to expunge the fixed effects (α_i) and uses OLS to estimate the parametric part of the regression equation. Afterwards $f(gini_{it})$ is estimated via a B-spline regression model. Moreover, we apply the semiparametric estimator by Robinson (1988) with the pooled data. Robinson's estimator, which was implemented in *Stata* by Verardi and Debarsy (2012), partials out the parametric part of the regression equation and runs kernel regressions on the residuals. As non-parametric estimations are sensible to outliers, we use the full set of control variables (and thus the narrower sample) for all semiparametric regressions.

[Figure 3 about here]

Figure 3 pictures the non-parametric part of these estimations, while the results for the linear part of the model are shown in Table 7 from the Appendix. The upper graph illustrates the estimated relationship from Robinson's semiparametric estimator, which we use with an Epanechnikov kernel function and cluster robust standard errors. In line with a corresponding pooled OLS estimation of a quadratic regression model (see, Table 7), Robinson's semiparametric estimator shows a hump-shaped relationship and a similar turning-point, lying roughly at a Gini of around 27.

The form of the relationship is less clear when it comes to the semiparametric fixed effects estimator. The lower part of Figure 3 indicates a concave relationship when the power of the B-splines is set to d(3). Yet in the default specification, with a power of d(4), the graph hints towards a 3rd order polynomial form, where saving tends to rise again at very high levels of inequality. A direct inclusion of a cubic term into a parametric fixed effects or pooled OLS model, however, yields no significant results (see, Table 7).[24]

In sum, semiparametric regressions yield no strong evidence against a quadratic functional form. As the simple fixed effects estimator is more efficient than the semiparametric alternatives, we regard this as sufficient evidence for a hump-shaped pattern between inequality and saving. Nonetheless, we will later discuss some arguments why saving rates may increase with inequality, when inequality is already very high.

5.3 Interactions with credit availability, financial development, and different time periods

5.3.1 Interactions with credit availability and financial development

Along the lines of previous studies (Smith, 2001, Alvarez-Cuadrado and El-Attar Vilalta, 2012) we suppose that the relation between inequality and saving may depend on the state of financial market development. The idea is that poorer households, who face a decline in relative income, need credit financing to keep up with rising consumption of the rich. Easy credit availability could thus be a precondition for expenditure cascades: In countries with liberalized financial markets expenditure cascades may dominate the link between inequality and saving, whereas Keynesian effects may prevail where credit financing is scarce.

To test for the presence of such a conditional effect we complement our baseline regression model with an interaction term, which is the product of inequality and a moderator variable measuring either credit availability or financial market liberalization:

$$\text{saving}_{it} = \alpha + \beta_1 \text{gini}_{it} + \beta_2 \text{credit}_{it} + \beta_3 \text{gini}_{it} \times \text{credit}_{it} + \beta' \text{X}_{it} + \alpha_i + \lambda_t + \epsilon_{it} \tag{3}$$

The first two Columns of Table 8 report the estimates for this interaction model (excluding and including the country fixed effects) with the ratio of private credit to GDP as the moderator variable (credit). Indeed, both the pooled OLS model of Column (1) and the fixed effects model of Column (2) yield strong evidence for a significant interaction effect. In both equations the product of Gini and credit is significantly negative, while the Gini has a significantly positive coefficient.

[Table 8 about here]

[24] As Baltagi and Li's (2002) estimator relies on a first difference transformation, while the standard fixed effects estimator is based on demeaning, results are not directly comparable.

Differentiating the equation in Column (2) with respect to inequality yields the marginal effect of inequality across different levels of credit, pictured as a downward sloping line in Figure 4. While the marginal effect of inequality on household saving is positive at low and average levels of credit, it becomes negative at a credit ratio of 130 percent. However, the surrounding 90% confidence intervals indicate that inequality exerts a significantly positive effect only with credit below 87% of GDP. Moreover, inequality only becomes significantly negative, when credit is above 165% of GDP, a threshold which for example the United States exceed since the early 2000s.

[Figure 4 about here]

A possible problem arising from the use of the credit ratio as an explanatory or moderator variable is that it could be endogenous with respect to the saving rate. To circumvent this problem, we employ the financial reform index, composed by Abiad et al. (2010), as a measure of credit market liberalization in Columns (3) and (4). Given that the financial reform index (*finreform*) is a de jure measure, it is free of endogeneity concerns. Yet, being based on sub-indices on subjects like capital account restrictions, interest rate controls, etc., the index is merely a rough proxy of credit availability.

When we substitute credit with *finreform* in the pooled OLS model of Column (3), the signs of the coefficients of inequality and the interaction term remain unchanged. Apparently, with highly regulated financial markets (low index values) a positive marginal effect of inequality prevails, but decreases and finally becomes negative with increasing financial liberalization (high index values). Nonetheless, in the fixed effects model of Column (4) the interaction effect is insignificant, which is not surprising given that most of the index variation stems from differences across countries.

As the effect of inequality depends on credit availability, it is questionable whether the concave and hump-shaped relationship is also robust to different states of financial development. To test for the presence of heterogeneous effects, we create a dummy variable, *credithigh*, which we set as 1 for values of credit to GDP above the sample median of 90%.[25] Then we effectively split our sample into a low-credit and a high-credit subsample by estimating the following model:

$$\text{saving}_{it} = \alpha + \beta_1 \text{gini}_{it} + \beta_2 \text{gini}_{it}^2 + (\beta_3 \text{gini}_{it} + \beta_4 \text{gini}_{it}^2 + \beta_5) \times \text{credithigh}_{it} + \beta' \text{X}_{it} + \alpha_i + \lambda_t + \epsilon_{it} \quad (4)$$

Column (5) of Table (8) reports our estimates for this model. The effect of inequality in

[25] The estimated coefficient of credithigh (-1.012) is insignificant (p-value: 0.131) in a model where *credithigh* serves as an additional regressor, but not as a moderator variable.

country-years with a low level of credit can be directly seized via β_1 and β_2, which indicate a significant hump-shaped relationship. At high-levels of credit $\beta_1+\beta_3$ and $\beta_2+\beta_4$ measure the inequality-saving relationship, indicating a concave relationship that is somewhat less pronounced than in the low-credit subsample.

[Figure 5 about here]

Based on the results from Column (5), Figure 5 illustrates the marginal effect of inequality at low and high levels of credit together with the 90% confidence intervals.[26] It shows that the Gini at which the marginal effect of inequality turns from positive to negative is somewhat higher in the low-credit group.[27] Moreover, very tight 90% confidence intervals in the low-credit subsample indicate that inequality exerts a significant positive effect on saving at a wider range of inequality values. Within the high-credit group inequality yields a significantly positive effect only at very low levels of inequality and becomes significantly negative for values of the Gini above 33.

Altogether, we find that the relation between inequality and saving tends to be positive with low credit availability and negative with high credit availability. Nonetheless, a hump-shaped relationship between inequality and saving prevails in both low and high-credit environments.

5.3.2 Inequality and saving after the financial crisis

Given that the risks of subprime lending to poorer households became obvious with the 2008-10 Global Financial Crisis (see, e.g. Rajan, 2010), the ability and willingness of poorer households to engage in expenditure cascades may have decreased. Thus the negative part of the hump-shaped relationship between inequality and saving may have vanished in the post-crisis period.

To test for this supposition we create a dummy variable (*postcrisis*), which we set as 1 for all observations after 2007. Then we effectively split our sample into a pre-crisis and a post-crisis subsample by estimating a nested regression model, similar to Equation 4.[28]

The results for $gini \times postcrisis$ and $gini^2 \times postcrisis$ in Column (6) of Table (8) indicate that the estimated coefficients of inequality have declined after the outbreak of the crisis. Yet the interacted terms are statistically insignificant and smaller than the main effects. Figure 6 plots the marginal effects of inequality received from the estimated equation. It shows that

[26] Generating this figure we benefited from the code provided by Arcand et al. (2015)

[27] The turning point is 31 in the low-credit group and 28 in the high-credit group.

[28] The most recent observations in our panel are from 2013, so that the post-crisis dummy marks all observations from the 2008 to 2013.

before 2008 inequality had a significantly positive effect on saving if the value of the Gini was below 25, a null effect at a Gini of around 30, and a significantly negative effect at Ginis above 35. In the post-crisis period, however, inequality never exerted a significant effect.

Summing up, the distinction between pre- and post-crisis periods confirms our supposition that a negative effect of inequality on household saving vanished in the wake of the financial crisis. Yet, the insignificance of the relationship between inequality and saving could as well be due to the loss of efficiency, given the small number of observations in the post-crisis period.

[Figure 6 about here]

5.4 Private saving, national saving, and the current account

Finally, we analyze whether the effect of inequality on household saving transmits to broader measures of saving and the current account balance. Although our theories of interest refer to household behavior, the household saving rate would be too narrow if richer households maintain a large volume of saving within incorporated enterprises. As it includes saving from both the household and the corporate sector, the use of private saving rates could thus be beneficial. Following previous cross-country studies on inequality and saving, we also look at national saving rates, which include saving by the government. National saving could be of interest as it measures the total amount of saving in the economy. Yet, its application is problematic if fiscal policy exerts offsetting effects.

Referring to studies that motivate our paper, we finally check whether the link between inequality and saving transmits to the current account. Being the balance between national saving and investment, we would expect that inequality has a similar influence on the current account as it has on saving.

[Table 9 about here]

Table 9 presents the results of regressions for these alternative dependent variables. To enhance comparability, each Column draws on a uniform sample of 517 observations. The regressors are identical to the small fixed effects model from our baseline table. Yet, as it is too closely related to public saving, which is part of the dependent variable, we drop the fiscal balance in the regressions for national saving.[29]

[29] One could as well argue that the fiscal balance is a direct component of the current account balance. Yet, because it is frequently used in the current account literature, we keep the fiscal balance as a regressor in Column (5).

As a benchmark reference, Column (1) repeats the baseline household saving regression, which is now based on the uniform sample. Column (2) reports results for the net private saving rate. Both Columns (3) and (4) cover national saving rates: In Column (3) national saving is measured net of fixed capital depreciation, in line with the concept that we adopt throughout this paper. Yet most previous studies use gross national saving rates, which we utilize in Column (4). Finally, Column (5) reports results for the current account balance. The Data Appendix describes the sources and derivations of the new dependent variables.

For each saving aggregate our results indicate a non-monotonic effect of inequality and the SLM-Test always confirms the existence of a hump-shaped relationship. Moreover, the shape of the relationship is always quite similar, with minor differences: For net private saving and even more so for net national saving, the effect of inequality appears to be positive at a wider range of Ginis. Yet, for gross national saving, the turning point of the hump-shaped relationship is again close to the peak value from the household saving regression. Even for the current account balance the effect of inequality is similar to the one we know from the household saving regressions. Apparently, the current account increases with rising inequality, if inequality is low, whereas it tends to decrease, when the Gini becomes larger than 30.[30]

In sum, the impact of inequality on household saving rates appears to transmit to broader saving aggregates. Moreover, although the drivers of current account balances are not the primary focus of this paper, our results also hint that inequality affects the current account in a non-monotonic way.

6 Discussion and conclusions

This paper shows that the marginal effect of inequality on saving is decreasing in the level of credit availability and financial liberalization. Above all, however, we find that the relationship between inequality and aggregate saving is hump-shaped, meaning that with higher levels of inequality the initially positive marginal effect of inequality decreases and eventually becomes negative.

An explanation for the decreasing marginal effect of inequality could be given by a non-linear adoption in household consumption behavior: If inequality only becomes gradually visible, the saving rates of poor and middle-class households possibly remain unchanged, while inequality is still rising from a low level. Thus aggregate saving would initially be dominated by an increasing income share of households with a high propensity to save. As

[30] The hump-shaped relationship also prevails with the full set of covariates and an unrestricted sample. Results are available on request.

inequality rises further, this positive effect on saving could be increasingly compensated by a changing behavior of households from the middle and lower ranks of the income distribution. When inequality becomes more and more visible, the incentive to engage in conspicuous consumption rises until the decrease in saving of poorer households dominates in aggregate.

Moreover, at high levels of inequality, further gains in inequality could increasingly result from a decline in the real income of poorer households. At some point, income may fall below a level that suffices to finance saving plus socially acceptable minimum consumption. Income losses will then be compensated by a reduction of the saving rate. When the latter starts to offset the direct effect from rising income concentration, the marginal effect of inequality on aggregate saving decreases and after some point becomes negative.

Our findings suggest that the inequality driven decrease in saving, at high levels of inequality, appears to have vanished since the outbreak of the global financial crisis. Even if inequality continues to rise, a permanent compensation of income losses via credit financing is hardly conceivable. Consequently, as soon as saving rates of low and middle income households have reached a floor at zero, it is likely that the Keynesian effect of a rising income concentration at the top will dominate.

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Appendix

Appendix A1: Data Description

Household saving rate (saving$_{hh}$): Is sourced from the OECD Economic Outlook (EO) and National Accounts (NA) Databases. The OECD calculates saving by subtracting household consumption expenditure from household disposable income plus the change in net equity of households in pension funds. Saving is reported as net of depreciation. The saving rate is calculated with saving in the numerator and the net household disposable income, plus the change in the net equity of households in pension funds, in the denominator. The formula for the saving rate in the System of National Accounts is (see, Catte and Boissinot, 2005):

$$s_t = S_t/YD_t = (YD_t - C_t)/YD_t = (r_t W_{t-1} + Y_t - T_t - C_t)/(r_t W_{t-1} + Y_t - T_t)$$

where s denotes the ratio of saving S, to disposable income, YD, and C is the value of consumption. Disposable income consist of labour income, Y, and capital income, rW, minus taxes and transfers, T.

Gross national saving (saving$_{gross}$): The World bank calculates gross saving (% of GDP) as gross national income less total consumption, plus net transfers.

Net national saving (saving$_{net}$): Net saving (% of GDP) is sourced from the OECD NA Database. It is defined as the difference between disposable income and final consumption expenditure plus an adjustment for the change in pension entitlements. Net saving is reported net of fixed capital depreciation.

Private saving (saving$_{prvt}$): Is calculated as net national saving less net saving of general government (% of GDP), which is also sourced from the OECD NA Database.

Current account: The current account balance (% of GDP) is sourced from the OECD EO database.

Gini coefficient (gini): Our preferred measures of income inequality is the Gini of net incomes from the Standardized World Income Inequality Database (SWIID, Version 5.0, released in October 2014) generated by Solt (2009, 2016). The SWIID is based on the UN World Income Inequality Database (WIID), and several other cross-country inequality datasets, data provided by national statistical offices and scholarly articles. As the source data is not directly comparable it is transformed and adjusted in several steps, described in Solt (2016). A very rough overview of the standardization procedure can be given as follows: 1. The data is sorted into categories by welfare definitions and by equivalence scale. Ginis of net and market inequality on the basis of household adult equivalent income from the Luxembourg Income Study (LIS) are added as a baseline, generating a dataset in which each country-year observation has data entries in at least one of thirteen categories. 3. Ratios between the variables in different categories are estimated as a function of country-decade, country, region and development status through various regression models. In further steps eleven series of estimates, comparable with the LIS net-income data, are calculated and combined into a single variable. 4. Possible measurement errors are corrected by using five-year weighted moving averages on all data points except those taken from the LIS and certain time periods.

$Gini_{LIS}$ and *ratio8020*: Our alternative measures of inequality come from the *Key Figures* of the Luxembourg Income Study. Both the Gini coefficient and the 80/20 percentile ratio rest on disposable household income.

Top 1% income share (top1inc): The data on top incomes is sourced from the World Top Incomes Database (WTID) by Atkinson et al. (2015). Whenever possible, we chose the standard series, which were given without any reference to divergent tax units or data sources. Exceptions are: The UK and Denmark where the data measures the income share of adults; and Finland where numbers are based on the income distribution survey (IDS). In contrast to our Gini data, top income shares are based on gross incomes before taxes and transfers. Series are mostly expressed as percentage of total income excluding capital gains, but there are also some differences in income definition, which we ignore in order to maintain a reasonably wide dataset. Due to these and other limitations on cross-country comparability, the inclusion of country fixed effects is crucial in regressions with top income shares.

Old-age dependency ratio (depend): Is obtained from the World Banks WDI database. It is defined as the share (in %) of the population aged 65 and older over the working-age population (people between 15 and 64).

Real growth rate of disposable income per capita (incgrow): Real household disposable income was taken from the OECD Economic Outlook and the OECD National Accounts Database. Growth rates reflect year on year variations.

Real interest rate (interest): The real interest rate is calculated by subtracting the inflation rate from the interest rate on long-term government bonds, which is sourced comes from the OECD Economic Outlook.

Fiscal balance (fiscal): The fiscal balance (cash surplus/deficit % of GDP) is collected from the WDI database, except when data from the OECD EO database is available (NLGQ: Government net lending, as a percentage of GDP)

Ln GDP per capita (ln(gdppc)): Is the natural logarithm of per capita GDP sourced from the OECD National Accounts Database. GDP is measured in US Dollars (000s) at constant prices and constant PPPs. The OECD base year is 2010.

Inflation rate (infl): The CPI inflation rate is sourced from the OECD Main Economic Indicators and expanded with World Bank WDI data.

Real house price development (houses): A quarterly index of real house prices is retrieved from the OECD Housing Prices database. We calculate the year on year growth rate on the annual averages of the quarterly data.

Real equity return (equities): Is measured as the yearly performance of the main national share price indices, sourced from the OECD Main Economic Indicators, minus the CPI inflation rate. According to the OECD data description the share indices are targeted to be national, all-share or broad, price indices.

Financial depth (credit and finreform): *Credit*, is the ratio (expressed in %) of private credit to GDP, from the WDI database. *Finreform* is the Financial Reform Index provided by Abiad et al. (2010). It is an institutional index that measures the overall level of financial market regulation on the basis of several sub-indices: Credit directions and requirements on central bank reserves, interest rate controls, entry barriers, aggregate credit ceilings, state ownership in the banking sector, capital account restrictions, prudential regulation and banking supervision, and securities market policies. The index ranges between 0 and 100. Higher values of the index imply more liberalized financial markets.

Countries in the sample are: Australia, Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Japan, Korea Rep., Luxembourg, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, United Kingdom, United States.

Appendix A2: Semiparametric and 3rd order polynomial regressions.

[Table 7 about here]

Figures

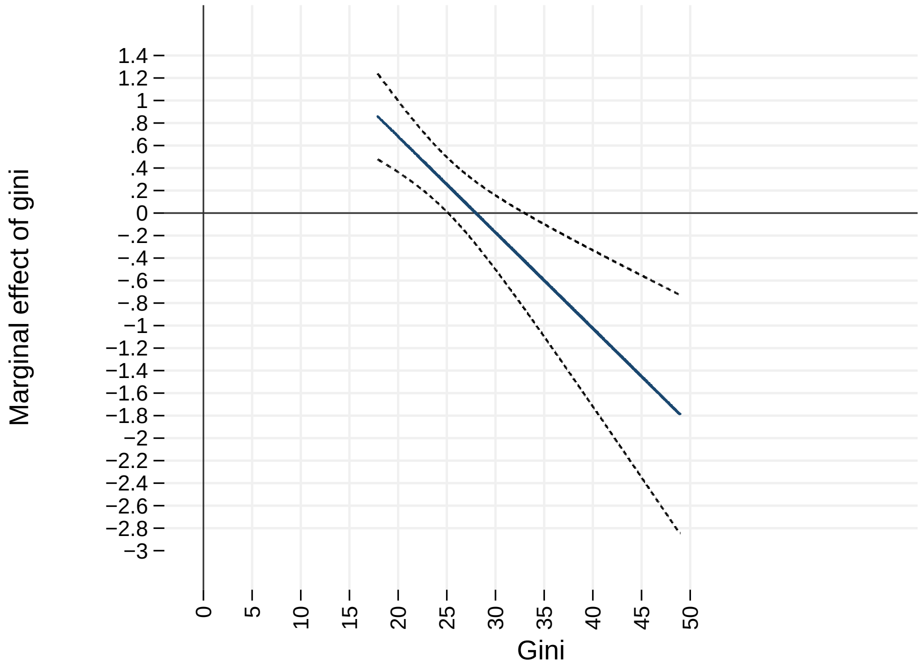

Figure 1 The marginal effect of inequality on saving at different levels of inequality. Values are calculated from the results of Table 2, Column (4).

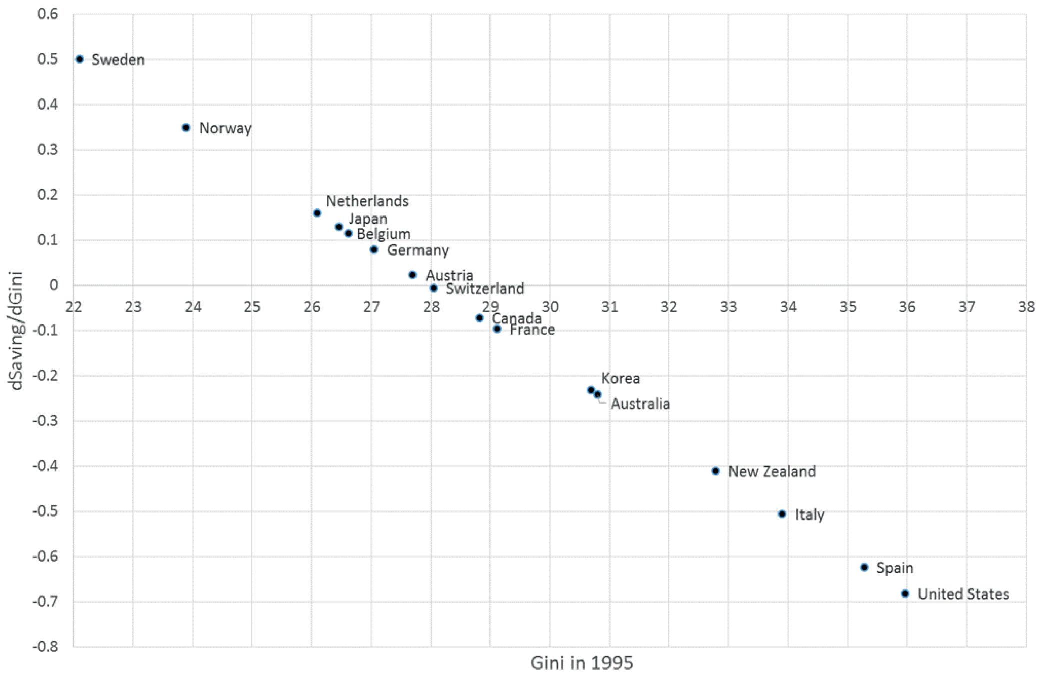

Figure 2 The marginal effect of inequality on saving at 1995 Gini levels. Values are calculated from the results of Table 2, Column (4).

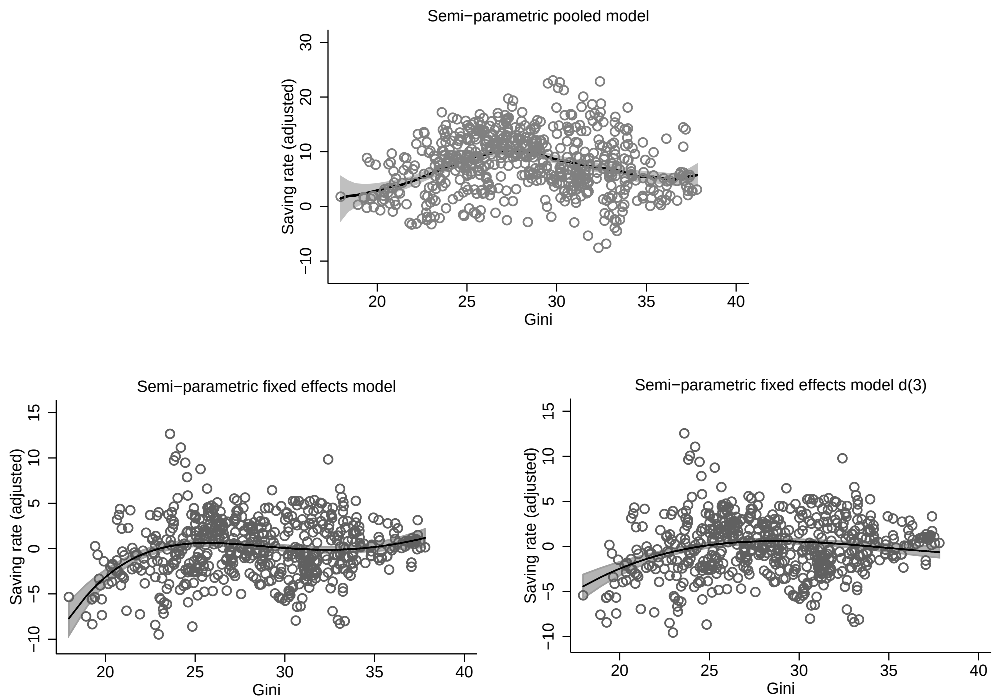

Figure 3 Partial fit of the relationship between saving and inequality. The points in each graph are partial residuals for the household saving rate; saving rates have been adjusted for the effects of the linear control variables (see, Eq. 2). Partial residuals of the fixed effects regressions are centered around the mean. Shaded areas correspond to 90% confidence intervals.

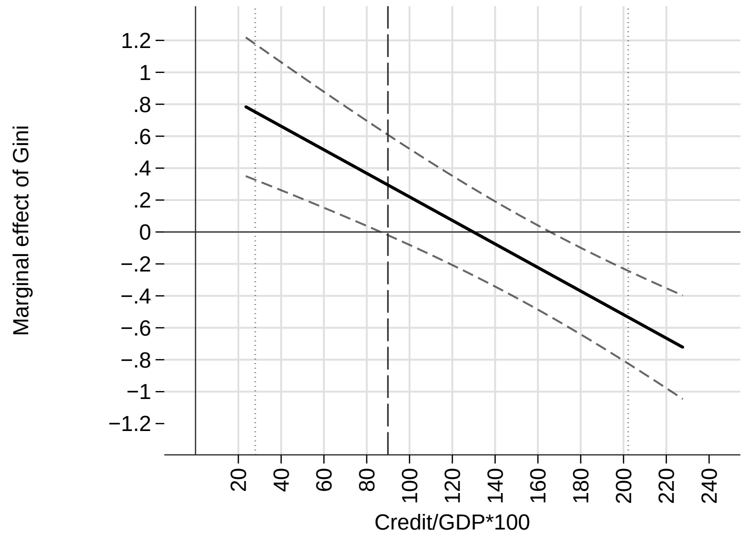

Figure 4 The marginal effect of inequality on saving across different levels of credit availability: Values are calculated using the results of Column (2) of Table (8). The downwards sloping line plots the marginal effect of inequality. Surrounding dashed lines represent the 90% confidence intervals. Vertical lines indicate the distribution of the credit to GDP ratio in the sample: dotted lines mark the first and 99th percentiles, the dashed line marks the median value.

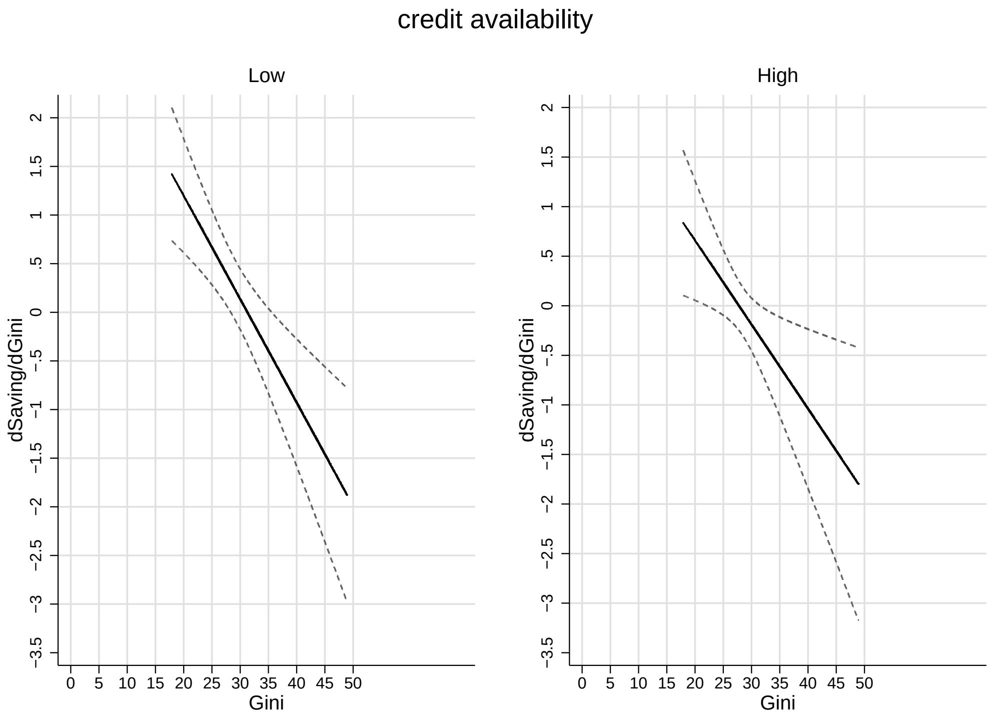

Figure 5 The marginal effect of inequality on saving with low and high credit availability (below and above 90% of GDP). Values are calculated using the results from Column (5) Table (8). The downwards sloping line plots the marginal effect of inequality at different levels of inequality. Surrounding dashed lines represent the 90% confidence intervals.

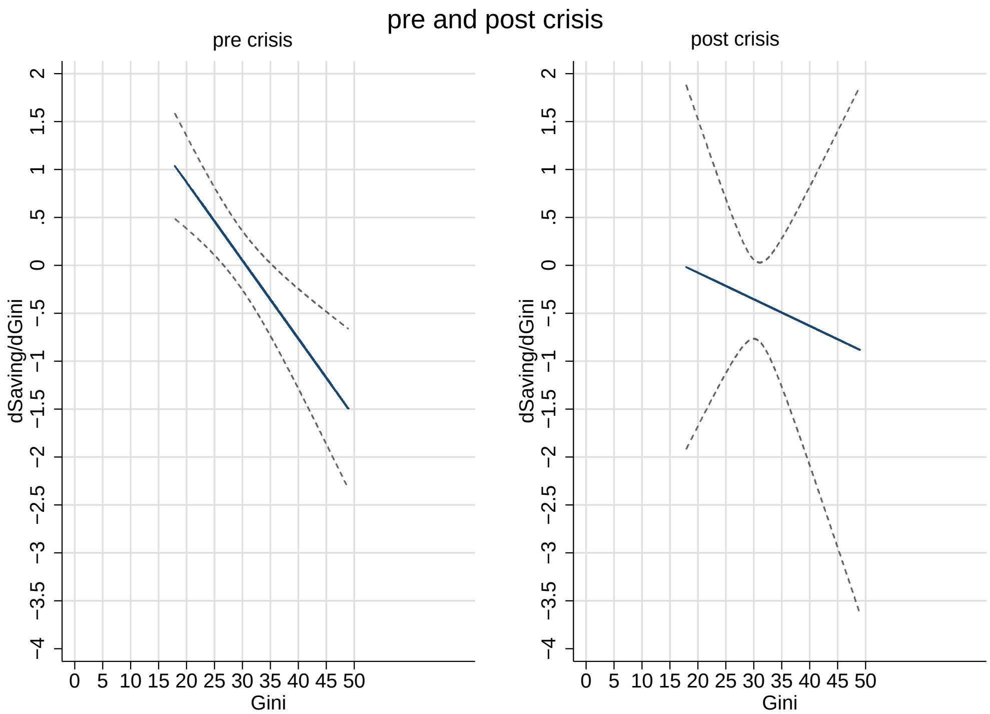

Figure 6 The marginal effect of inequality on saving before and after the Global Financial Crisis (before and after 2008). Values are calculated using the results from Column (6) Table (8). The downwards sloping line plots the marginal effect of inequality at different levels of inequality. Surrounding dashed lines represent the 90% confidence intervals.

Tables

Table 1 Summary statistics

Variable	Mean	Std. Dev.	Min.	Max.	N
saving_{hh}	7.930	5.989	-9.043	25.776	792
gini	28.282	4.403	17.964	48.74	792
gini_{LIS}	28.328	4.123	19.7	37.1	142
ratio8020	2.313	0.335	1.759	3.12	142
top1inc	8.01	2.679	3.97	18.33	427
depend	20.888	4.732	6.433	36.018	792
incgrow	2.391	2.958	-11.046	15.995	792
interest	3.045	2.935	-14.992	20.998	792
fiscal	-2.397	4.593	-32.554	18.696	792
ln(gdppc)	10.303	0.329	8.762	11.346	766
infl	3.986	3.602	-4.48	24.54	792
equities	4.337	23.498	-47.79	105.33	749
houses	1.656	7.092	-17.241	38.831	685
credit	89.863	44.031	20.84	227.753	757
finreform	75.086	23.418	9.524	100	527
saving_{prvt}	7.875	4.047	-4.215	23.285	527
saving_{net}	7.996	5.783	-12.653	31.164	713
saving_{gross}	24.272	5.408	6.118	41.745	723
current account	-0.145	4.639	-14.575	16.232	766

Table 2 Baseline regression models

	(1) POLS	(2) POLS	(3) FE	(4) FE	(5) FE	(6) FE	(7) FE	(8) FE
gini	-0.117	1.427	-0.0373	2.386***	0.138	3.218***	0.0513	3.159***
	(0.168)	(1.079)	(0.195)	(0.659)	(0.204)	(0.635)	(0.160)	(0.711)
$gini^2$		-0.0263		-0.0426***		-0.0533***		-0.0537***
		(0.0182)		(0.0130)		(0.0107)		(0.0119)
depend	-0.325**	-0.309**	-0.671***	-0.677***	-0.676***	-0.715***	-0.849***	-0.882***
	(0.143)	(0.140)	(0.118)	(0.128)	(0.188)	(0.156)	(0.203)	(0.152)
incgrow	0.448***	0.468***	0.271***	0.274***	0.385***	0.382***	0.426***	0.417***
	(0.101)	(0.107)	(0.0582)	(0.0611)	(0.0680)	(0.0642)	(0.0583)	(0.0494)
interest	-0.192	-0.175	-0.100	-0.0965	-0.130	-0.103	0.0306	-0.0215
	(0.158)	(0.157)	(0.108)	(0.112)	(0.176)	(0.156)	(0.191)	(0.159)
fiscal	-0.475***	-0.463***	-0.416***	-0.436***	-0.422***	-0.433***	-0.349**	-0.363***
	(0.155)	(0.155)	(0.103)	(0.0964)	(0.111)	(0.0959)	(0.127)	(0.111)
ln(gdppc)					-0.346	-1.114	-5.421	-7.502
					(4.933)	(4.035)	(8.714)	(7.654)
infl					0.0588	0.0887	0.196	0.201
					(0.162)	(0.104)	(0.220)	(0.168)
equities					-0.0155**	-0.0134**	-0.0151	-0.0163
					(0.00663)	(0.00513)	(0.0118)	(0.00994)
houses					-0.0758***	-0.0714**	-0.0539**	-0.0450*
					(0.0262)	(0.0271)	(0.0226)	(0.0236)
credit					-0.0249	-0.0152	-0.0370	-0.0252
					(0.0215)	(0.0181)	(0.0232)	(0.0202)
year-dummies	No	No	No	No	No	No	Yes	Yes
Observations	792	792	792	792	616	616	616	616
Countries	29	29	29	29	27	27	27	27
R-sq	0.223	0.239	0.433	0.458	0.549	0.583	0.573	0.608
Turning-Point		27.09		27.97		30.20		29.40
CI 90%				[25.04; 33.23]		[26.91; 34.02]		[26.59; 32.32]
Slope: $gini_{min}$		.48		.85***		1.30***		1.23***
Slope: $gini_{max}$		-1.15*		-1.79***		-2.00***		-2.11***
SLM p-val		0.14		.005		.0002		.0003

Notes: Table reports pooled OLS (POLS) and fixed-effects (FE) regressions. Dependent variable is the saving rate of the household sector. Cluster robust standard errors are reported in parentheses. The bottom part of the table reports the turning points of the inequality effect and the results of the Sasabuchi-Lind-Mehlum (SLM) test for a hump-shaped relationship. CI-90% denotes the 90% Fieller confidence intervals for the turning point. To ease comparison slopes at $gini_{min}$ and $gini_{max}$ are uniformly measured at the bounds of the maximum sample of 792 observations, i.e. at Ginis of 18 and 49. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3 Lagged regressors, IV, and System-GMM

	(1) Lag T-1 FE	(2) 2-SLS FE	(3) System GMM	(4) System GMM	(5) System GMM
$saving_{t-1}$			0.946***	0.945***	0.906***
			(0.0344)	(0.0343)	(0.0315)
gini	2.190***	2.526***	1.032**	0.714	0.658
	(0.608)	(0.662)	(0.516)	(0.487)	(0.475)
$gini^2$	-0.0393***	-0.0442***	-0.0176*	-0.0125	-0.0125
	(0.0120)	(0.0125)	(0.00955)	(0.00885)	(0.00828)
depend	-0.689***	-0.741***	0.00848	-0.00434	0.0189
	(0.138)	(0.142)	(0.0427)	(0.0442)	(0.0302)
incgrow	0.252***	0.232	0.136	0.101	0.445***
	(0.0895)	(0.316)	(0.0998)	(0.0930)	(0.0906)
interest	-0.136	-0.116	-0.0442	-0.0485	0.153**
	(0.122)	(0.152)	(0.0459)	(0.0672)	(0.0683)
fiscal	-0.326***	-0.453***	-0.00852	0.0162	-0.0430
	(0.0945)	(0.154)	(0.0384)	(0.0375)	(0.0450)
ln(gdppc)					2.825***
					(0.799)
infl					0.221***
					(0.0809)
equities					-0.0213*
					(0.0115)
houses					-0.0997***
					(0.0211)
credit					-0.000831
					(0.00589)
year-dummies	No	No	No	Yes	Yes
Observations	793	765	789	789	616
Countries	29	29	29	29	27
R-sq	0.418	0.453			
Hansen p-val		0.889	0.0856	1.000	1.000
AR(1) p-val			0.0000358	0.0000363	0.000658
AR(2) p-val			0.816	0.937	0.220
Instruments		9	20	72	77
Turning-Point	27.86	28.58	29.32	28.54	26.37
SLM p-val		.0006	.0419	.0865	.111

Notes: Dependent variable is the saving rate of the household sector. Column (1) reports a fixed effects model with cluster robust standard errors. All regressors are lagged for one period. Column (2) reports a 2-SLS regression where incgrow, interest and fiscal are instrumented with L(1/2). Columns (3)-(5) report one-step system GMM estimations with cluster robust standard errors, a collapsed instrument matrix and orthogonal deviations. All variables except depend are treated as endogenous. Instruments are the second and third lag of the explanatory variables in levels for the transformed equation and the first lag in differences for the level equation. The null of the Hansen J-test is the validity of all instruments. AR(1) p-val and AR(2) p-val report the p-values of the Arellano-Bond test for autocorrelation of the residuals. The bottom part of the table reports the turning points of the inequality effect and the results of the Sasabuchi-Lind-Mehlum (SLM) test for a hump-shaped relationship. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 4 Multiple-imputation estimates

	(1) POLS	(2) FE	(3) FE	(4) FE
gini_{mi}	1.388	2.179***	2.939***	2.917***
	(1.042)	(0.665)	(0.646)	(0.696)
$\text{gini}^2{}_{mi}$	-0.0256	-0.0389***	-0.0487***	-0.0496***
	(0.0176)	(0.0130)	(0.0112)	(0.0119)
depend	-0.310**	-0.683***	-0.712***	-0.883***
	(0.140)	(0.126)	(0.159)	(0.156)
incgrow	0.468***	0.274***	0.384***	0.419***
	(0.108)	(0.0612)	(0.0657)	(0.0504)
interest	-0.176	-0.0975	-0.108	-0.0173
	(0.157)	(0.112)	(0.156)	(0.161)
fiscal	-0.463***	-0.435***	-0.431***	-0.361***
	(0.155)	(0.0972)	(0.0983)	(0.113)
ln(gdppc)			-1.059	-7.421
			(4.105)	(7.754)
infl			0.0829	0.201
			(0.109)	(0.171)
equities			-0.0136**	-0.0161
			(0.00526)	(0.0101)
houses			-0.0726**	-0.0459*
			(0.0279)	(0.0238)
credit			-0.0159	-0.0261
			(0.0183)	(0.0204)
year-dummies	No	No	No	Yes
Observations	792	792	616	616
Countries	29	29	27	27
Turning-Point	27.06	28.07	30.19	29.4

Notes: Table presents multiple-imputation estimates of the baseline pooled OLS (POLS) and fixed effects (FE) regression models. Dependent variable is the saving rate of the household sector. Cluster robust standard errors are reported in parentheses. The final line of the table reports the turning points of the inequality effect. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 5 Alternative Data

	(1) POLS	(2) FE	(3) FE	(4) FE	(5) FE	(6) FE
gini_{LIS}	3.870**	3.310**				
	(1.406)	(1.236)				
gini_{LIS}^2	-0.0675***	-0.0574**				
	(0.0231)	(0.0216)				
ratio8020			50.45**			
			(23.43)			
ratio8020^2			-9.837*			
			(4.915)			
top1inc				-0.355	0.557	-0.125
				(0.343)	(1.328)	(0.419)
top1inc^2					-0.0421	
					(0.0502)	
gini						2.572**
						(1.057)
gini^2						-0.0470**
						(0.0222)
depend	0.000483	-0.627**	-0.650**	-0.771***	-0.805***	-0.809***
	(0.168)	(0.225)	(0.237)	(0.131)	(0.143)	(0.101)
incgrow	0.214	0.344***	0.363***	0.276***	0.257***	0.217***
	(0.166)	(0.0835)	(0.0898)	(0.0638)	(0.0664)	(0.0674)
interest	0.380	-0.115	-0.108	-0.257**	-0.269**	-0.273**
	(0.343)	(0.124)	(0.119)	(0.107)	(0.105)	(0.0985)
fiscal	-0.381***	-0.440***	-0.458***	-0.465***	-0.505***	-0.513***
	(0.117)	(0.0600)	(0.0722)	(0.0990)	(0.0910)	(0.0915)
Observations	143	143	143	430	430	427
Countries	25	25	25	18	18	18
R-sq	0.194	0.476	0.464	0.602	0.608	0.637
Turning-Point	28.67	28.83	2.56			27.38
CI 90%	[24.90; 30.79]	[25.73; 32.63]	[2.17; 4.14]			[24.57; 45.7]
Slope at gini_{min}	1.44**	1.24***				.88***
Slope at gini_{max}	-2.74***	-2.32***				-2.03**
SLM p-val	.0111	.00885	.0342			.0456

Notes: Table reports pooled OLS (POLS) and fixed-effects (FE) regressions. Dependent variable is the saving rate of the household sector. Cluster robust standard errors are reported in parentheses. The bottom part of the table reports the turning points of the inequality effect and the results of the Sasabuchi-Lind-Mehlum (SLM) test for a hump-shaped relationship. CI-90% denotes the 90% Fieller confidence intervals for the turning point. To ease comparison slopes at gini_{min} and gini_{max} are measured at the bounds of the baseline sample, i.e. at Ginis of 18 and 49. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 6 Restricted country or time samples

	(1) Excluding outliers	(2) G7 countries	(3) 1961-1995	(4) 1980-2013	(5) 1990-2013	(6) 2000-2013
gini	3.300**	1.415	1.809**	4.110***	3.321***	0.530
	(1.404)	(1.149)	(0.821)	(0.851)	(1.062)	(1.314)
$gini^2$	-0.0565**	-0.0242	-0.0324**	-0.0654***	-0.0496**	-0.00379
	(0.0255)	(0.0182)	(0.0148)	(0.0139)	(0.0186)	(0.0248)
depend	-0.829***	-0.691***	-0.235	-0.801***	-0.740***	-0.265
	(0.150)	(0.154)	(0.193)	(0.163)	(0.200)	(0.157)
incgrow	0.275***	0.364**	0.231***	0.361***	0.318***	0.332***
	(0.0468)	(0.105)	(0.0762)	(0.0543)	(0.0495)	(0.0462)
interest	-0.250	0.0203	-0.376***	-0.154	-0.108	-0.0599
	(0.180)	(0.146)	(0.0900)	(0.156)	(0.165)	(0.100)
fiscal	-0.462***	-0.474***	-0.507***	-0.388***	-0.277***	-0.274***
	(0.113)	(0.115)	(0.0847)	(0.0899)	(0.0951)	(0.0567)
ln(gdppc)	5.622*	4.769		-3.878	-7.876	4.857
	(3.186)	(2.757)		(5.320)	(5.655)	(5.426)
infl	0.0581	0.309***		0.0373	-0.0533	-0.431***
	(0.126)	(0.0823)		(0.148)	(0.220)	(0.146)
equities	-0.0116*	-0.0136**		-0.0167**	-0.0144**	-0.0238***
	(0.00615)	(0.00447)		(0.00640)	(0.00694)	(0.00613)
houses	-0.0700	-0.0427**		-0.0751**	-0.0840*	-0.0677
	(0.0443)	(0.0149)		(0.0305)	(0.0474)	(0.0420)
credit	-0.0389**	-0.0500**		-0.00940	0.00173	-0.0126
	(0.0177)	(0.0190)		(0.0170)	(0.0141)	(0.0148)
Observations	497	212	354	560	451	291
Countries	26	7	18	27	27	27
R-sq	0.513	0.846	0.335	0.560	0.496	0.380
Period	1971-2013	1971-2013	1961-1995	1980-2013	1990-2013	2000-2013
Turning Point	29.21	29.22	27.94	31.4	33.47	-
SLM p-val	.0285	.171	.0261	.0004	.034	-

Notes: Table reports fixed-effects (within) regressions with cluster robust standard errors in parentheses. Dependent variable is the saving rate of the household sector. The bottom part of the table reports the turning points of the inequality effect and the p-values from the Sasabuchi-Lind-Mehlum (SLM) test for a hump-shaped relationship. Column (1) drops the top and bottom 5% of saving and inequality; (2) includes only the G7 economies; (3)-(6) draw on different time periods. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 7 Semiparametric and 3rd order polynomial regressions

	(1) Semi-POLS	(2) Semi-FE	(3) Semi-FE	(4) POLS	(5) POLS	(6) FE
gini				4.805***	16.54	6.082
				(1.097)	(10.79)	(6.440)
$gini^2$				-0.0849***	-0.502	-0.157
				(0.0192)	(0.389)	(0.234)
$gini^3$					0.00486	0.00123
					(0.00458)	(0.00275)
depend	-0.127	0.0489	0.0405	-0.168	-0.133	-0.687***
	(0.169)	(0.171)	(0.172)	(0.154)	(0.157)	(0.186)
incgrow	0.645***	0.339***	0.338***	0.659***	0.658***	0.384***
	(0.117)	(0.0277)	(0.0279)	(0.123)	(0.120)	(0.0641)
interest	-0.0500	0.0650	0.0690	-0.137	-0.0590	-0.0874
	(0.207)	(0.0720)	(0.0739)	(0.190)	(0.210)	(0.154)
fiscal	-0.449***	-0.231***	-0.229***	-0.496***	-0.471***	-0.425***
	(0.160)	(0.0633)	(0.0627)	(0.155)	(0.156)	(0.0943)
ln(gdppc)	-0.587	-13.12***	-13.08***	0.205	-0.169	-1.246
	(2.978)	(3.063)	(3.041)	(2.889)	(2.998)	(4.119)
infl	0.512**	0.149	0.151	0.470**	0.543***	0.105
	(0.206)	(0.0916)	(0.0924)	(0.186)	(0.191)	(0.115)
equities	0.0176*	-0.00472*	-0.00447*	0.0183*	0.0188*	-0.0131**
	(0.0102)	(0.00261)	(0.00258)	(0.0103)	(0.0105)	(0.00497)
houses	-0.113*	-0.0457***	-0.0463***	-0.120*	-0.118*	-0.0717**
	(0.0618)	(0.0144)	(0.0143)	(0.0597)	(0.0615)	(0.0275)
credit	-0.00952	0.00169	0.00195	-0.0106	-0.00805	-0.0148
	(0.0178)	(0.00662)	(0.00657)	(0.0178)	(0.0176)	(0.0179)
Observations	616	581	581	616	616	616
Countries	27	27	27	27	27	27
R-sq	0.338	0.408	0.408	0.386	0.391	0.582

Notes: Dependent variable is the saving rate of the household sector. Column (1)-(3) report the linear part of semiparametric models, estimated with Robinson's semiparametric regression estimator (Column 1) and the Baltagi and Li fixed effects estimator (Columns 2 and 3). Columns (4)-(6) report pooled OLS (POLS) and fixed-effects (FE) regressions with cluster robust standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 8 Interactions

	(1) POLS	(2) FE	(3) POLS	(4) FE	(5) FE	(6) FE
gini	1.146***	0.958***	1.332***	0.390	3.325***	2.499***
	(0.322)	(0.295)	(0.371)	(0.338)	(0.971)	(0.732)
$gini^2$					-0.0532***	-0.0408***
					(0.0163)	(0.0120)
credit	0.311***	0.185***				-0.0195
	(0.0910)	(0.0459)				(0.0163)
ginixcredit	-0.0116***	-0.0074***				
	(0.00297)	(0.00159)				
finreform			0.351*	0.157		
			(0.170)	(0.109)		
ginixfinreform			-0.0161***	-0.00614		
			(0.00543)	(0.00378)		
credithigh					17.71	
					(21.90)	
ginixcredithigh					-0.968	
					(1.544)	
$gini^2$xcredithigh					0.0107	
					(0.0268)	
postcrisis						37.64
						(40.64)
ginixpostcrisis						-2.020
						(2.616)
$gini^2$xpostcrisis						0.0269
						(0.0417)
Observations	616	616	451	451	648	616
Countries	27	27	21	21	27	27
R-sq	0.430	0.612	0.509	0.539	0.595	0.609

Notes: Table reports pooled OLS (POLS) and fixed-effects (FE) regressions with cluster robust standard errors in parentheses. Dependent variable is the saving rate of the household sector. Control variables (depend, incgrow, interest, fiscal, equities, houses, ln(gdppc), infl) are omitted for clarity.* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 9 Alternative dependent variables

	(1) Household-Saving	(2) Private-Saving	(3) National-Saving (net)	(4) National-Saving (gross)	(5) Current Account
gini	3.057***	2.755**	3.086***	1.894***	2.602***
	(0.865)	(1.060)	(0.843)	(0.494)	(0.709)
gini^2	-0.0547***	-0.0461**	-0.0497***	-0.0333***	-0.0436***
	(0.0123)	(0.0167)	(0.0138)	(0.00811)	(0.0123)
depend	-0.814***	-0.351*	-0.650***	-0.277*	0.131
	(0.279)	(0.185)	(0.164)	(0.142)	(0.122)
incgrow	0.296***	0.316***	0.470***	0.247***	-0.140*
	(0.0622)	(0.0854)	(0.0644)	(0.0830)	(0.0711)
interest	-0.0405	-0.0522	-0.282***	-0.339***	-0.00218
	(0.179)	(0.109)	(0.0938)	(0.115)	(0.120)
fiscal	-0.524***	-0.360***			0.0112
	(0.0964)	(0.0779)			(0.0665)
Observations	517	517	517	517	517
Countries	25	25	25	25	25
R-sq	0.439	0.310	0.419	0.238	0.0939
Turning Point	27.93	29.88	31.06	28.43	29.83
CI 90%	[22.68; 30.91]	[24.91; 33.03]	[28.18; 34.60]	[24.56; 31.89]	[26.93; 33.75]
Slope: gini_{min}	1.09**	1.1**	1.3***	.69***	1.03***
Slope: gini_{max}	-2.31***	-1.76***	-1.78***	-1.37***	-1.67***
SLM p-val	.0104	.0148	.0018	.0025	.0022

Notes: Table reports fixed-effects (within) regressions with cluster robust standard errors in parentheses. The bottom part of the table reports the turning points of the inequality effect and the results of the Sasabuchi-Lind-Mehlum (SLM) test for a hump-shaped relationship. CI-90% denotes the 90% Fieller confidence intervals for the turning point. Slopes at gini_{min} and gini_{max} are uniformly measured at ginis of 18 and 49.* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$