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Ups and Downs in Finance, Ups without Downs in Inequality

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Abstract

The upswing in finance over the past several decades has led to rising inequality, but do downswings in finance lead to a symmetric decline in inequality? In this paper, we analyze the asymmetry of the effect of ups and downs in financial markets, as well as the effect of increased capital requirements and the bonus cap on national earnings inequality. We use administrative employer–employee linked data on earnings from 1990 to 2017 for twelve countries. Additionally, we use data on earnings from bank reports, from 2009 to 2017 in thirteen European countries. We find a strong asymmetry in the effects of financial ups and downs on earnings inequality, a mitigating effect of rising capital requirements on the contribution of finance to inequality, and a restructuring effect of the bonus cap for the earnings of financiers, while neither policy affects absolute levels of earnings inequality.

Keywords: inequality, finance, financial crisis, regulation

JEL classification: N2 financial markets and institutions; D31 personal income, wealth and their distributions; G38 government policy and regulation

Résumé

La hausse de la finance au cours des dernières décennies a entraîné une hausse des inégalités, mais les ralentissements de la finance entraînent-ils une baisse symétrique des inégalités? Dans cet article, nous examinons l'asymétrie de l'effet des hausses et des ralentissements des marchés financiers, ainsi que l'effet de l'augmentation des exigences en matière de capital et du plafonnement des primes sur l'inégalité des salaires nationaux. Nous utilisons des données administratives couplées employeur-employé sur les salaires de 1990 à 2017 pour douze pays. De plus, nous employons des données sur les salaires provenant des rapports bancaires, de 2009 à 2017, dans 13 pays européens. Nous constatons une forte asymétrie dans les effets des hausses et des ralentissements financières sur l'inégalité des salaires, un effet de mitigation de l'augmentation des exigences de capitalisation sur la contribution de la finance à l'inégalité, et un effet de restructuration du plafonnement des primes pour les salaires des financiers, alors qu'aucune des deux mesures n'affecte les niveaux absolus d'inégalité des salaires.

Mots clés: inégalité, finance, crise financière, régulation

Contents

Ups and Downs in Finance, Ups without Downs in Inequality

1 Introduction

The financial crisis sparked a double-barreled public debate about the role of finance in contemporary inequality. For one, the left-leaning #OccupyWallStreet movement put the blame on finance for elevating economic inequality to Gilded Age levels (Calhoun 2013). Conversely, many public intellectuals indicted finance not so much for engendering inequality in times of economic upswing, but for not taking responsibility in times of downswing and crisis (Rajan 2008; Attali 2009). The latter commentators took umbrage that many financial firms continued to pay large bonuses while they defaulted or were bailed out by their respective governments. The case of AIG even inspired draft US legislation to impose a prohibition of bonus payments in government-backed banks, although the bill was never passed (Thomas 2009).

Since then, research concentrating on the period of upswing in finance in the 1990s and 2000s has shown that financialization contributed significantly to rising income inequality (Tomaskovic-Devey and Lin 2011; Lin and Tomaskovic-Devey 2013; Godechot 2012; Kus 2012; Dünhaupt 2014; Flaherty 2015; Denk and Cournède 2015; Roberts and Kwon 2017; Huber, Petrova, and Stephens 2020; Lin and Tobias Neely 2020; see also Hager 2020). Systematic comparisons with cross-country macro data suggest that an increasing number of handsomely paid financiers is the key determinant of increasing top income shares (Kus 2012; Dünhaupt 2014; Godechot 2016; Huber et al. 2020). However, while the first prong of the public controversy has been scrutinized in depth, the second has received little attention and the finance–inequality nexus in the post-crisis period has been left unexplored. It remains obscure whether downswings in finance have resulted in decreased earnings in the financial industry and thus downswings in national inequality.

After 2008, two developments in finance reversed its contribution to the upswing in inequality: downward trends in financial activity, and financial regulation. In the wake of the financial crisis, many economic actors reduced or altered their engagement in financial markets. As a result, the volume of activity on financial markets declined and the profitability of financial firms plummeted. Secondly, governments, central banks and the international financial regulatory community sought to limit the excessive risk-taking of banks and individual bankers. Hence, policies imposing higher capital requirements have been widely implemented and the EU introduced the so-called bonus cap in 2013, which aims to limit the risk-taking of individual financiers.

We would like to thank Louison Carroué and Julie Fournier for excellent research assistance in collecting bank reports. We would like to thank the participants of the SocioCongress 2021 in Geneva for helpful feedback.

Although both the decline in financial activity and policy interventions have no immediate impact on financiers' earnings, they could contribute to a decline in financial sector earnings. Earnings in finance are intended to be tied to performance through bonuses, and banks use various unit-level formulas to calculate bonus pools (Godechot 2017). Hence, a decline in profitability in a particular market should translate into a decline in earnings in that banking unit. Moreover, a reduction in bank leverage dilutes the profitability of more risk-exposed activities and should therefore depress bonus pools further. Finally, a bonus cap inflates wage costs because fixed wages are less downwardly flexible than bonuses in a financial downturn. Rational banks should therefore negotiate a reduction in total remuneration, and rational financiers should accept such a reduction as they may be better off with a less volatile remuneration structure.

However, the asymmetric assignment of responsibility for profits and losses to financiers could limit the impact of a decline in bank profitability on earnings (Godechot 2017). Additionally, previous research on the "75 percent tax" for millionaires in France (Guillot 2021) shows that top earners such as financiers have considerable bargaining power to circumvent the costs of policies that target their earnings. What is more, Murphy (2013) predicted that the European bonus cap would increase financiers' fixed remuneration to match the remuneration in unregulated sectors, harm the profitability of the European banking sector, and incentivize financiers to take more risk (Murphy 2013). Hence, the same organizational processes could underpin the response to both market shocks and regulatory pressure. The joint analysis of the effects on inequality of market-led and policy-led disruptions in finance promotes a better understanding of the mechanisms behind persistent earnings inequality. If neither market processes nor regulation can reverse financial excesses and their inequitable consequences, we should take heed of the risk of letting finance take hold.

To assess the contribution of finance to earnings inequality before and after the financial crisis and to measure the effects of financial regulation, we use two novel and unique databases. First, building on a large research consortium, we rely on a composite of administrative employer–employee linked data on earnings from 1990 to 2017 for twelve countries (Canada, Denmark, Norway, Sweden, France, Germany, Netherlands, Spain, Czechia, Hungary, South Korea, and Japan), further complemented with World Bank indicators of financialization (GFDD) and income share estimates from Piketty and Saez (2003)[1] for the United States.[2] This allows us to document the evolution of top earnings shares and their distribution between financial and nonfinancial sectors on a much larger scale than has previously been done (Godechot 2012; Bakija, Cole, and Heim 2012; Bell and Van Reenan 2014). In addition, this allows us to measure the asymmetry of the inequality effect of financial market activities and the introduction of stricter capital requirements. Second, to analyze the impact of the European bonus cap, we use a self-collected dataset on financiers' remuneration from Bank reports, from

1 Updated to 2018 by Emmanuel Saez in 2020 (https://eml.berkeley.edu/~saez/TabFig2018.xls).

2 We complement these estimates with estimates based on Current Population Survey (CPS) data.

2009 to 2017 in thirteen European countries (Germany, Austria, Belgium, Denmark, Finland, France, Greece, Ireland, Italy, Netherlands, Spain, Sweden, United Kingdom), which we complement with data on balance sheet indicators from COMPUSTAT.

Based on this array of evidence, we make four contributions. First, we corroborate and substantiate previous research that shows the dramatic contribution of financiers' earnings to the upswing in inequality. Second, we demonstrate that the contribution of finance to inequality did not decrease with the decline in financial market activity. Third, we show that post-crisis increases in capital requirements have at best yielded a very modest reduction in the contribution of finance to inequality. Finally, we show that the bonus cap led to a short-run decline in variable remuneration and a long-run increase in fixed remuneration, leaving total remuneration, and thus the role of the financial sector in income inequality, unchanged.

The remainder of the paper proceeds in four parts. In the first section, we describe the data used. In the second section, we delineate the trends in top earnings shares and the overrepresentation of financiers in the top 1 percent earnings share and analyze the effects of financial activity on their evolution. We devote the third section to measuring the effects of financial regulation. We, first, analyze the impact of increasing bank capital on earnings inequality and, second, evaluate the adjustment of European banks' remuneration practices to the introduction of the bonus cap. In the conclusion, we discuss explanations of the asymmetry between upswing and downswing effects and discuss policy implications of our research.

2 Data

Administrative employer–employee linked data

Building on a large-scale collaboration, we use administrative employer–employee linked data for twelve countries: Canada, Denmark, Sweden, Norway, France, Germany, Netherlands, Spain, Czechia, Hungary, South Korea, and Japan (cf. Table A1). Data on earnings inequality complement those typically used for income inequality but offer the distinct advantage of capturing inequality at the workplace.

We base our analysis on one billion worker-year observations and up to 150 million worker observations per year. Some countries, such as Canada, Denmark, Norway, Sweden, Netherlands, and France, provide exhaustive information on the working population and permit very reliable estimates for small groups in small units. In other countries (Germany, Spain, South Korea, and Japan), we use samples of between 4 and 8 percent of the working population. With regard to the common socioeconomic research, the latter samples are very large and enable reliable estimates of top earning shares. However, the estimates of earnings shares of financiers in top fractiles can be more fragile. This is especially the case in Germany, where, in addition, top earnings are top coded around the top decile threshold. To overcome this limitation, we imputed top earnings in Germany.[3] For the United States we used secondhand estimates from Piketty and Saez (2003)[4] of national inequality based on US Social Security data. These estimates do not contain a decomposition of top shares between financiers and non-financiers.

With this large-scale research endeavor we can track the evolution of top earnings shares, and their decomposition between finance (insurance sector included) and non-finance in a variety of political economies (Esping-Andersen 1990; Hall and Soskice 2001): two "liberal" Northern American economies (Canada, USA), three Scandinavian "social-democratic" economies (Denmark, Norway, Sweden), three "corporatist" western European economies (France, Germany, Netherlands), two eastern European transitioning economies (Czechia, Hungary) and one "Southern Europe Economy" (Spain, cf. Katrougalos 1996), and two East Asian economies (South Korea, Japan). This diversity ensures that our results are not conditional on a specific institutional setting.

World Bank GFDD database

To measure the effects of post-crisis capital regulations we use the World Bank Global Financial Development Database (GFDD) from 1988 to 2017. This database provides common indicators of financialization (Kus 2012; Godechot 2016) such as *Stock market total value traded to GDP* (series GFDD.DM.02) or *Stock market capitalization to GDP* (GFDD.DM.01). It further contains aggregated indicators of banks' profits (such as *Bank return on equity pretax* – GFDD.EI.10) and indicators of bank capital, that is, firstly, *bank capital to total assets* (GFDD.SI.03) and, secondly, *bank regulatory capital to risk-weighted assets* (GFDD.SI.05). The latter ratio is a good indicator of capital regulation, as it corresponds well with the benchmarks used to set capital requirements for banks in Europe and the United States. However, this indicator is based on bank-internal risk measurement, and banks have a certain de facto freedom to apply their own risk-weighting strategies. We therefore additionally use the indicator for bank capital to total assets, which is more robust to endogenous processes within banks. We further use two indicators to capture the effect of the leverage requirements, which is the second pillar of capital regulation introduced in the countries analyzed.

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3 In Germany, our imputation strategy uses contemporaneous and lagged information from both individuals and workplaces to predict high earnings, using a tobit function estimated for multiple education by gender for east/west German populations. Code and further discussion available upon request.

4 Updated to 2018 by Emmanuel Saez in 2020 (https://eml.berkeley.edu/~saez/TabFig2018.xls).

European bank reports

With the introduction of the Capital Requirements Directive III (CRD III) in 2009, banks were required to disclose the remuneration of a part of their employees in their annual reports (CEBS 2010). We exploit these annual bank reports to create our novel dataset on the remuneration of financiers for twenty-five European banks from 2009 to 2017 (see Table A2). The data form an unbalanced panel as directive CRD III was not implemented concurrently in every member state.[5] We include four main variables from these reports in our analysis: fixed remuneration, variable remuneration, total remuneration, and the number of "material risk takers" subject to the disclosure requirements. Variable remuneration consists of both deferred and non-deferred payments in the form of monetary payments, or instruments such as shares. Fixed remuneration is a fixed annual payment in monetary form. Total remuneration is the sum of fixed and variable remuneration. The bank reports contain aggregate statistics on total remuneration and the number of incumbents for "material risk takers." These are provided both at the bank level and for specific functions or units, such as managing directors, investment banking, retail banking, and independent control functions. Banks have the task of identifying these material risk takers that are defined as employees "whose professional activities have a material impact on the institution's risk profile" (CEBS 2010). This somewhat vague definition refers to employees in executive positions, independent control functions and decision makers in the banking units.

We use the COMPUSTAT database to complement our dataset on financiers' remuneration with several balance sheet indicators for respective banks. We use indicators for total assets and number of employees to capture the size of the balance sheets and workforce of banks. As a measure of bank profitability, we use the indicator earnings before interest, taxes, depreciation and amortization (EBITDA), which is the sum of sales revenue less cost of goods sold and selling, general and administrative expenses.

3 The contribution of financiers' earnings to inequality and its asymmetry in upswings and downswings

To begin with, we analyze the contribution of finance to earnings inequality. Previous literature examined this relationship broadly through two channels, the financialization of non-financial firms or the increasing number of well-paid bankers. The first strand of literature argues that with financialization, non-financial firms prioritize shareholder interest through dividend disbursement and downsizing of the workforce (Goldstein 2012; Dünhaupt 2014; Huber et al. 2020; Hager 2020). This structural change

5 We include only those banks in our models for which data is available since 2013 in order not to distort regression parameters by increasing sample size post-regulation.

in non-financial firms and performance-based executive compensation widens the pay gap between top managers and labor and thus increases earnings inequality. Moreover, this literature suggests that financialization contributes to a weakening of labor institutions by promoting labor flexibility, thereby contributing to greater earnings inequality (Volscho and Kelly 2012; Jung 2015; Flaherty 2015; Darcillon 2015).

A second stream of literature focuses on the mechanisms of inequality aggravation at the micro level. With financialization, the financial labor market developed conducive conditions and a disproportionate increase in financiers' wages, thereby yielding a significant financial wage premium (Tomaskovic-Devey and Lin 2011; Godechot 2012; Philippon and Reshef 2012; Bakija, Cole, and Heim 2012; Denk and Cournède 2015; Bell and Van Reenen 2014; Godechot 2017).

In this section, we present strong evidence that the rising number of well-paid bankers is the key mechanism through which finance contributes to rising earnings inequality. With our administrative earnings data, we produce estimates analogous to those compiled by Piketty and Saez (2003)[6] for the United States, for twelve additional countries. Figure 1 displays the evolution of top 1 percent earnings shares in thirteen high wage countries and shows the familiar figure of strong rising inequalities during the 1990s and 2000s. It further shows that the financial crisis of 2008 led to a significant drop in the top 1 percent's earnings share in most countries. This drop was temporary and lasted only one or two years. However, in some cases, for example in Canada or France, it spanned up to 2013. In some countries (USA, Canada, Germany, Sweden) the dot. com crisis led to a similar temporary drop in the top 1 percent earnings share.

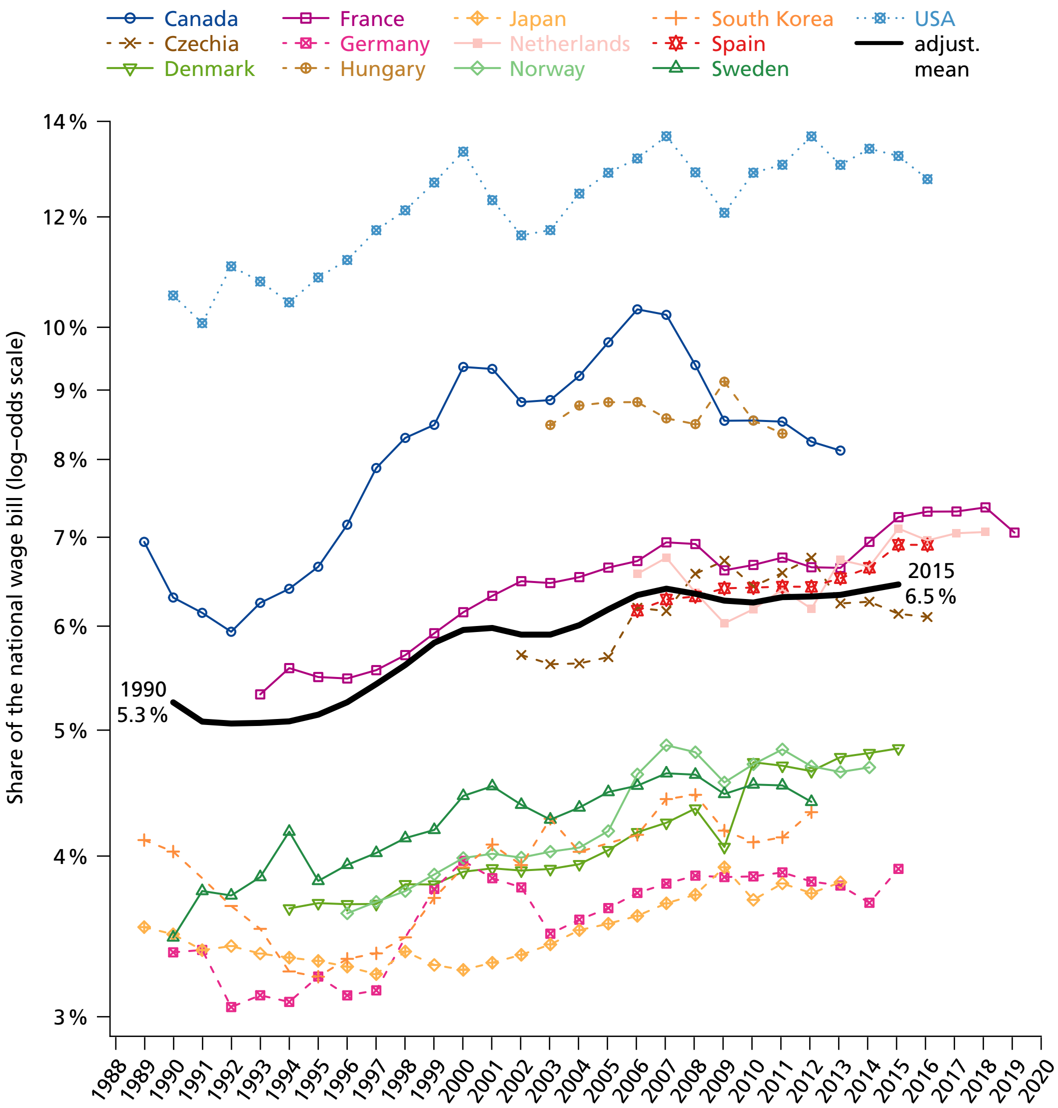
To measure finance's contribution to the evolution of inequality, we first use a simple additive decomposition method already present in the literature (Bakija, Cole, and Heim 2012; Godechot 2012; Philippon and Reshef 2012; Bell and Van Reenen 2014). We decompose the evolution of the top earnings share as the sum of the evolution of two shares: that of top earners who work in finance and that of top earners who work outside finance (equation 1).

$$\Delta S_{nat\_top1\%} = \Delta S_{nat\_top1\%\ \&\ finance} + \Delta S_{nat\_top1\%\ \&\ non\text{-}finance} \quad (1)$$

with earnings share $S_{nat\_top1\%\ \&\ sector} = \Sigma_i (w_{i,sector} > P99_{nat}) / \Sigma_i w_i$
w_i: wage of individual i,
$P99_{nat}$: P99 threshold of the national wage distribution
sector: either finance or non-finance

6 Updated to 2018 by Emmanuel Saez in 2020 (https://eml.berkeley.edu/~saez/TabFig2018.xls).

Figure 1 Evolution of top 1 percent wage earnings in thirteen countries



Note: For the USA, we use Saez's updated estimates of Piketty and Saez (2003; cf. Table B5 in https://eml.berkeley.edu/~saez/TabFig2018.xls). For other countries source description, cf. Appendix A1.

In a second step, we express the contribution of finance to the evolution of inequality as the ratio of the evolution of the earnings share of workers belonging both to the finance sector and to the national top 1 percent over the evolution of the earnings share of the national top 1 percent (equation 2).

$$Finance\_Contribution = \Delta S_{nat\_top1\%\ \&\ finance} / \Delta S_{nat\_top1\%} \quad (2)$$

For instance, in Table 1, the Canadian top 1 percent share moved from 6.0 percent to 10.4 percent between 1992 and 2006, which is a 0.31-point increase per year. During the same period, the earnings share of the Canadian financiers of the top 1 percent moved from 0.9 percent to 2.2 percent, which is a 0.09-point increase per year. Therefore, the contribution of finance to inequality is 29 percent (0.09/0.31).

In Table 1, we first apply this formula to the pre-crisis upswing. To produce both more reliable and more conservative estimates of finance's contribution to inequality, we apply this formula to inequality waves of maximum amplitude. We thus use the inequality minima as the start date and the inequality maxima as the end date for each country during the upswing period.

This exercise confirms that finance contributed massively to the rise in inequality, as has been shown for the United States, the United Kingdom and France (Bakija, Cole, and Heim 2012; Philippon and Reshef 2012; Godechot 2012; Bell and Van Reenen 2014). For instance, in Sweden, half of the increase in the top 1 percent share went to employees in the financial sector, as did 61 percent in South Korea, 43 percent in France, 40 percent in Hungary, 39 percent in Denmark, and 29 percent in Germany, Canada, and Norway. We find two exceptions, the Czechia, where finance does not contribute to the modest rise in inequality at all, and Japan, where finance was negatively correlated with rising inequality. This latter finding is likely a result of the severe banking crisis in Japan in the 1990s (Hoshi and Kashyap 1999).

In contrast with the upswinging inequality period, finance does not make a homogeneous contribution to the post-crisis downswing in inequality. In three countries (Norway, France, and Japan) finance has a negative contribution to the drop in inequality; that is, finance still contributes to increasing inequality while national inequality decreases. In addition, in three countries (Canada, Denmark, and the Netherlands), finance makes a much smaller contribution to the decrease in inequality than to the increase. Finally, finance's contribution to the decline is greater than its contribution to the increase in four countries (Sweden, Germany, Hungary, and South Korea). Note here that the decline in inequality is short-lived in most countries in our sample. However, these results suggest that in most countries, the contribution of finance to inequality is larger than its contribution to the decrease.

This previous decomposition has the advantage of being additive. However, the results depend on the size of the financial sector and are measurable only in the case of a substantial change in top earnings share. Therefore, we consider a second indicator of finance's contribution to inequality which neutralizes the differences in size between countries with a large (for example, Canada) or a small (for example, Czechia) financial sector. For this, we estimate an over-representation ratio (OR) of the earnings of the financial sector in the national top 1 percent with an odds ratio of proportions:

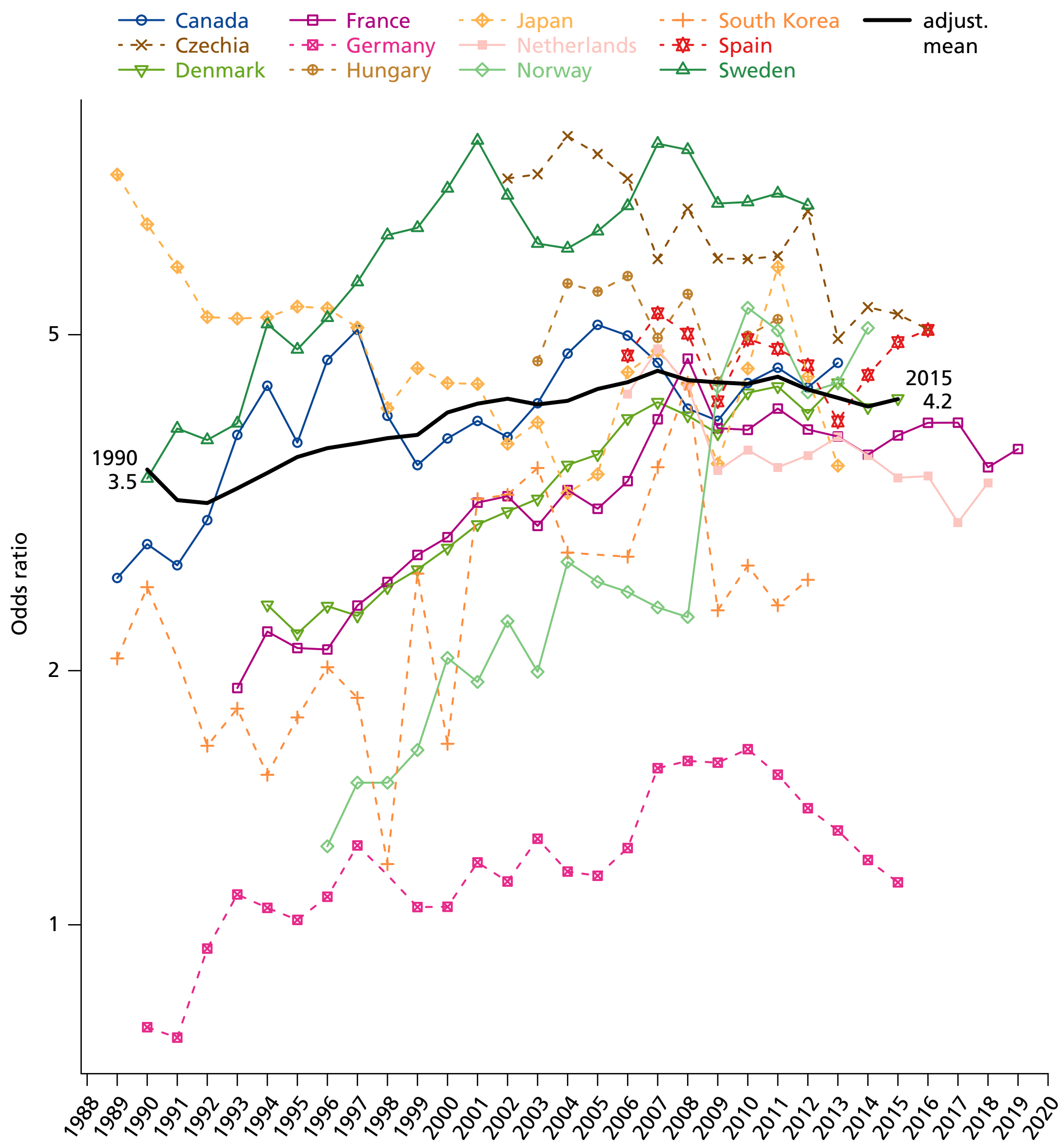
$$OR_{top1\%\ \&\ finance} = \\ = [S_{top1\%\ \&\ finance} / (1-S_{top1\%\ \&\ finance})] / [S_{top1\%\ \&\ non\ finance} / (1-S_{top1\%\ \&\ non\ finance})] \quad (3)$$

Table 1 Finance contribution to the pre-financial crisis upswing in inequality and to the financial crisis drop in inequality

Country	Pre-financial crisis inequality upswing					Financial crisis inequality downswing			
	Years	Bottom top 1%	Peak top 1%	Increase	Finance contribution	Years	Drop top 1%	Decrease	Finance contribution
Germany	1992–2008	3.1%	3.9%	+0.05%	19%	2008–2014	3.7%	–0.03%	47%
Denmark	1994–2008	3.7%	4.4%	+0.05%	39%	2008–2009	4.1%	–0.29%	15%
Japan	1997–2009	3.3%	4.0%	+0.06%	–27%	2009–2010	3.7%	–0.22%	–53%
Sweden	1990–2007	3.5%	4.7%	+0.07%	50%	2007–2009	4.5%	–0.08%	70%
Spain*	2006–2008	6.2%	6.4%	+0.08%	112%				
South Korea	1995–2008	3.3%	4.5%	+0.10%	61%	2008–2010	4.1%	–0.18%	141%
Hungary	2003–2009	8.5%	9.2%	+0.11%	40%	2009–2011	8.4%	–0.38%	49%
Norway	1996–2007	3.6%	4.9%	+0.12%	19%	2007–2009	4.6%	–0.16%	–25%
France	1993–2007	5.4%	7.0%	+0.12%	43%	2007–2009	6.7%	–0.16%	–25%
Czechia	2003–2009	5.7%	6.8%	+0.19%	0%	2009–2010	6.5%	–0.29%	7%
Netherlands	2006–2007	6.6%	6.8%	+0.19%	152%	2007–2009	6.1%	–0.36%	60%
Canada	1992–2006	6.0%	10.4%	+0.31%	29%	2006–2013	8.2%	–0.31%	20%
USA (CPS)	1992–2002	6.6%	9.6%	+0.30%	19%				
USA** (Bakija, Cole, and Heim 2012)	1993–2005	12.7%	17.0%	+0.35%	29%				
UK (Bell and Van Reenen 2014)	1999–2008	7.1%	8.9%	+0.20%	78%				

Note: The share of the national top 1% in Canada increased from 6.0% to 10.4% of the national wage bill during the inequality upswing between 1992 and 2006, an increase of +0.31 percentage points per year. 29% of this increase went to members of the national top 1% working in finance. During the post-financial crisis drop in inequality, the top 1% share dropped to 8.2% of the national wage bill. The top 1% share decreased at a yearly rate of –0.31%. 20% of this decrease was for members of the national top 1% working in finance.
* We could not detect a drop in inequality with the financial crisis in the Spanish data (2006–2016).
** Estimates from Bakija, Cole, and Heim are based on a notion of income rather than a notion of earnings.

Figure 2 Overrepresentation of financiers' earnings in the national top 1 percent



With this indicator, we calculate the overrepresentation of financiers' earnings in the national top 1 percent in contrast to that of workers working in the rest of the economy.

In Figure 2, we plot this indicator of finance's contribution to inequality during the downswing period. On average in 1992, finance earnings were 3.2 more represented in the national top 1 percent than nonfinancial earnings. Finance's overrepresentation moves to 4.5 in 2007 and drops to 4.2 in 2015 and the growth rate was three times the fall rate.

Less finance, less inequality? The asymmetry of the redistribution of earnings through financialization

With the global financial crisis, the financial sector's profitability declined consistently. For instance, as shown in Figure A1 in the appendices, the average pretax return on equity for banks (ROE) was well above 10 percent, with a peak of 19 percent in 2005. In contrast, the ROE after 2007 fell consistently below 10 percent, with a historical low of 7 percent in 2009. Nevertheless, despite dismal returns, financial earnings remained high and contributed to the preservation of high levels of inequality.

To further analyze the asymmetry of the financial sector's influence on inequality in upswing and downswing periods, we can draw on financial market indicators that are strongly associated with rising inequality in the existing literature. These include stock market volume and capitalization to GDP (Godechot 2012 and 2016; Kus 2012; Dünhaupt 2014; Huber, Petrova, and Stephens 2020). Several aspects justify the use of such indicators. First, previous research showed that the 1980–2007 upswing in finance is much more an upswing of financial market–related activities than global growth of the financial sector (Greenwood and Scharfstein 2013). Second, because banks set their bonus pools based on the profits of their internal units rather than the bank's overall profits (cf. Godechot 2017), indicators that capture trading opportunities are more likely to affect top financiers' remuneration than indicators focused on the bank's overall profits.

Figure 3 for volume of trade on national stock exchanges and Figure A2 in the appendices for capitalization as a share of GDP show a similar pattern, followed by most, if not all, countries in our sample: a sharp increase in trade volume between 1992 and 2001, followed by a decline between 2001 and 2003, a renewed increase between 2003 and 2008, which reverses between 2008 and 2013.

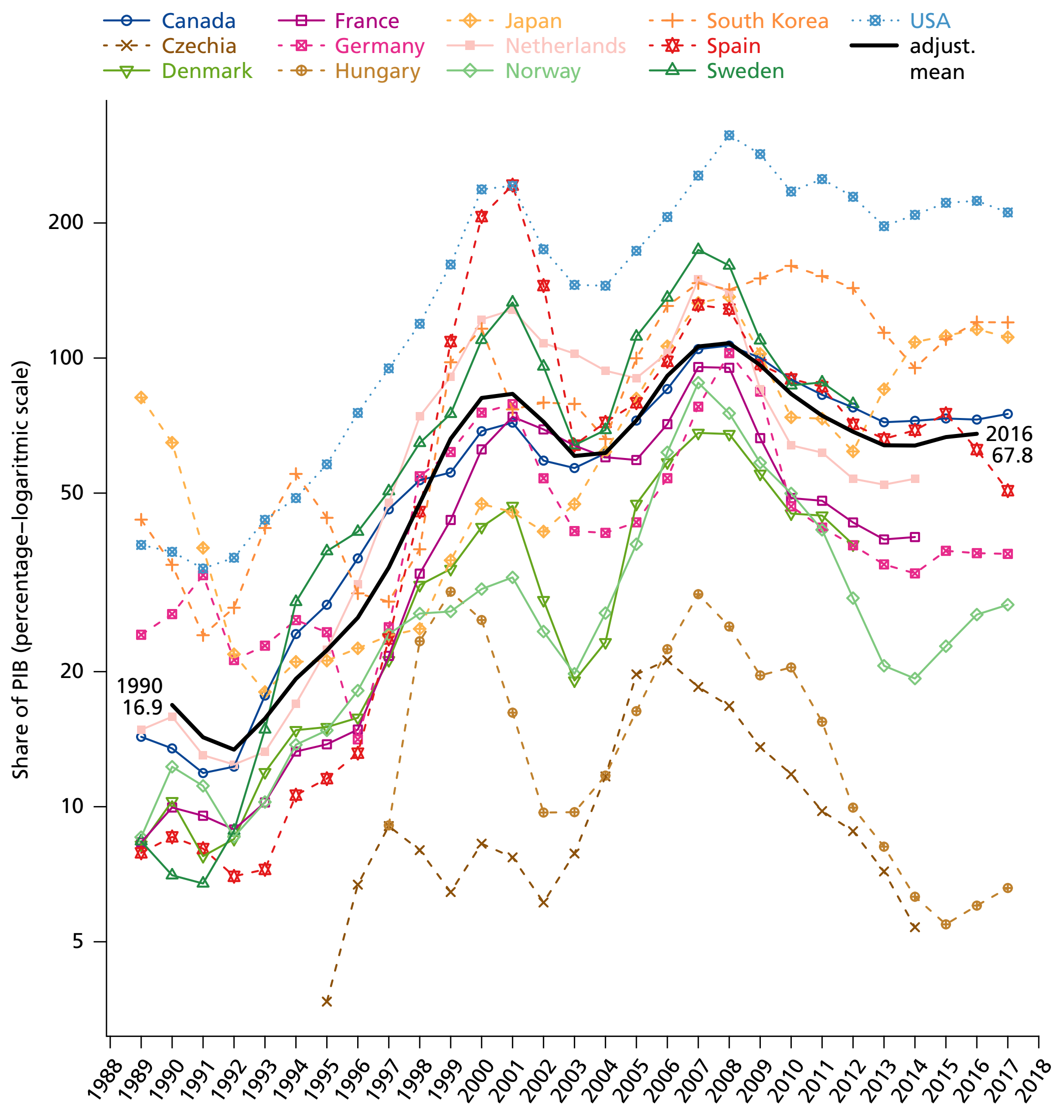
In Table 2, we follow Godechot (2016) and estimate the impact of trading volume on the top 1 percent's share and finance earnings' overrepresentation in top 1 percent, respectively, with a simple panel model that employs country and year fixed effects.

$$ineq_{ctry,t} = volume_{ctry,t} + controls_{ctry,t} + country_{ctry} + year_t + u \quad (4)$$

Therefore, we measure the impact of the country-specific evolution in trading volume on country-specific measures of inequality. We introduce the same control variables as in Godechot (2016), namely GDP per capita, union rate, and importation rate.[7] All independent variables are lagged. We provide two different measures of finance's activity asymmetry. First, we focus specifically on crisis years and thus interact finance activity

7 It is worth noting that union density does not have a negative impact on inequality, as in previous work (Kristal 2010; Godechot 2016). Several differences might explain these discrepancies. We focus here on top earning shares instead of top income shares. Our regression focuses on recent years in which the impact of unions may have changed.

Figure 3 Volumes of trade on national stock exchanges



with a dummy variable capturing the periods of decrease in financial activities in 2001–2003 and 2008–2013. Alternatively, to take into account all country-specific negative shocks, we introduce the cumulative sum of negative shocks as a supplementary variable: $(\sum_t[(\Delta_{t,t\text{-}1}volume_{ctry})^*(\Delta_{t,t\text{-}1}volume_{ctry}<0)])$. This variable captures the marginal effect of falls in financial activity on inequality.[8]

Models 1 and 4 show that the indicator of volume of stocks traded has a strong impact on inequality and an even stronger impact on finance's contribution to inequality. A one standard deviation increase in stocks traded augments the top 1 percent earnings share

8 An independent variable x_{jt} can be decomposed as:
$x_{jt} = \sum_t[(\Delta_{t,t\text{-}1}x_j)^*(\Delta_{t,t\text{-}1}x_j<0)] + \sum_t[(\Delta_{t,t\text{-}1}x_j)^*(\Delta_{t,t\text{-}1}x_j>=0)]+x_{j0}$

Table 2 The asymmetric impact of trading volume on inequality and financial earnings' overrepresentation in top earnings share

	Top 1% share			Finance earnings' over-representation in top 1%		
	(1)	(2)	(3)	(4)	(5)	(6)
GDP per capita	0.20***	0.18***	0.21***	0.89***	0.88***	0.92***
	(0.06)	(0.05)	(0.05)	(0.11)	(0.10)	(0.11)
Union rate	0.21***	0.18***	0.21***	0.19**	0.16**	0.39***
	(0.04)	(0.04)	(0.04)	(0.08)	(0.06)	(0.07)
Importation rate	0.07	0.07	0.07	0.16	0.15	0.19**
	(0.08)	(0.07)	(0.08)	(0.11)	(0.11)	(0.09)
Volume of stocks traded/GDP	0.23**	0.53***	0.23**	0.77***	1.32***	0.94***
	(0.09)	(0.11)	(0.10)	(0.23)	(0.31)	(0.21)
Volume of stocks traded/GDP × Years in (2001–2003, 2008–2013)		–0.59***			–0.95**	
		(0.16)			(0.34)	
Cumulative sum of drops in volume of stocks traded/GDP			–0.01			–0.75***
			(0.10)			(0.14)
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
R^2 (full model)	0.61	0.64	0.61	0.49	0.52	0.54
R^2 (proj model)	0.11	0.16	0.11	0.32	0.36	0.38
Num. obs.	250	250	250	223	223	223
Num. groups: country	13	13	13	12	12	12

Note: OLS models with country and year fixed effects and panel corrected standard errors in parenthesis. All independent variables are one-year lagged. Dependent and independent variables are country-demeaned and standardized. Hence, one standard deviation of stock exchange volume increases by 0.23 standard deviation the top 1% earnings share. ***p<0.01; **p<0.05; *p<0.1

by 0.23 standard deviation and the overrepresentation of finance in the top 1 percent by 0.77 standard deviation. Models 2, 3, 5, and 6 show the asymmetry in the effect of this indicator in times of downswing. In times of upswing, a one standard deviation increase in trading volume produces, respectively, a 0.5 standard deviation increase in inequality and a 1.3 standard deviation in finance's contribution to inequality. During the downswing, the marginal impact of decreasing trading volume is significantly lower than its positive effect in upswing periods (–0.6 and –1.0). Overall, the main effect for the share of the top 1 percent is eliminated and that for the financial sector's contribution to inequality is divided by 4. When we use the cumulative drop variable as an alternative measure of asymmetric causality, we find no significant difference between the decrease and the increase in trading volume in terms of the change in the share of the top 1 percent.[9] However, we do find a clear asymmetric effect of trading volume on our indicator of finance's relative contribution to inequality: when trading volume increases by one standard deviation, finance earnings' overrepresentation in the top 1 percent increases by 0.94 standard deviation. Conversely, when trading volume decreases by one standard deviation, the decreasing effect is only 0.19 (that is, 0.94–0.75). Finally, Table A3 shows a similar significant asymmetrical impact of upswings and downswings when using capitalization to GDP as the indicator of financial market activity.

9 Note, however, that the asymmetric effect is clearly significant when we use capitalization to GDP as the financialization indicator. Cf. Table A3.

Thus, these results show that while upswings in finance strongly contribute to upswings in inequality, downswings in finance do not contribute significantly to downswings in inequality. Following this analysis of the asymmetric effects of financial market activities, we examine the effects of financial regulation aimed at reducing excessive risk-taking in the financial industry and, consequently or as an intermediate step, financial earnings.

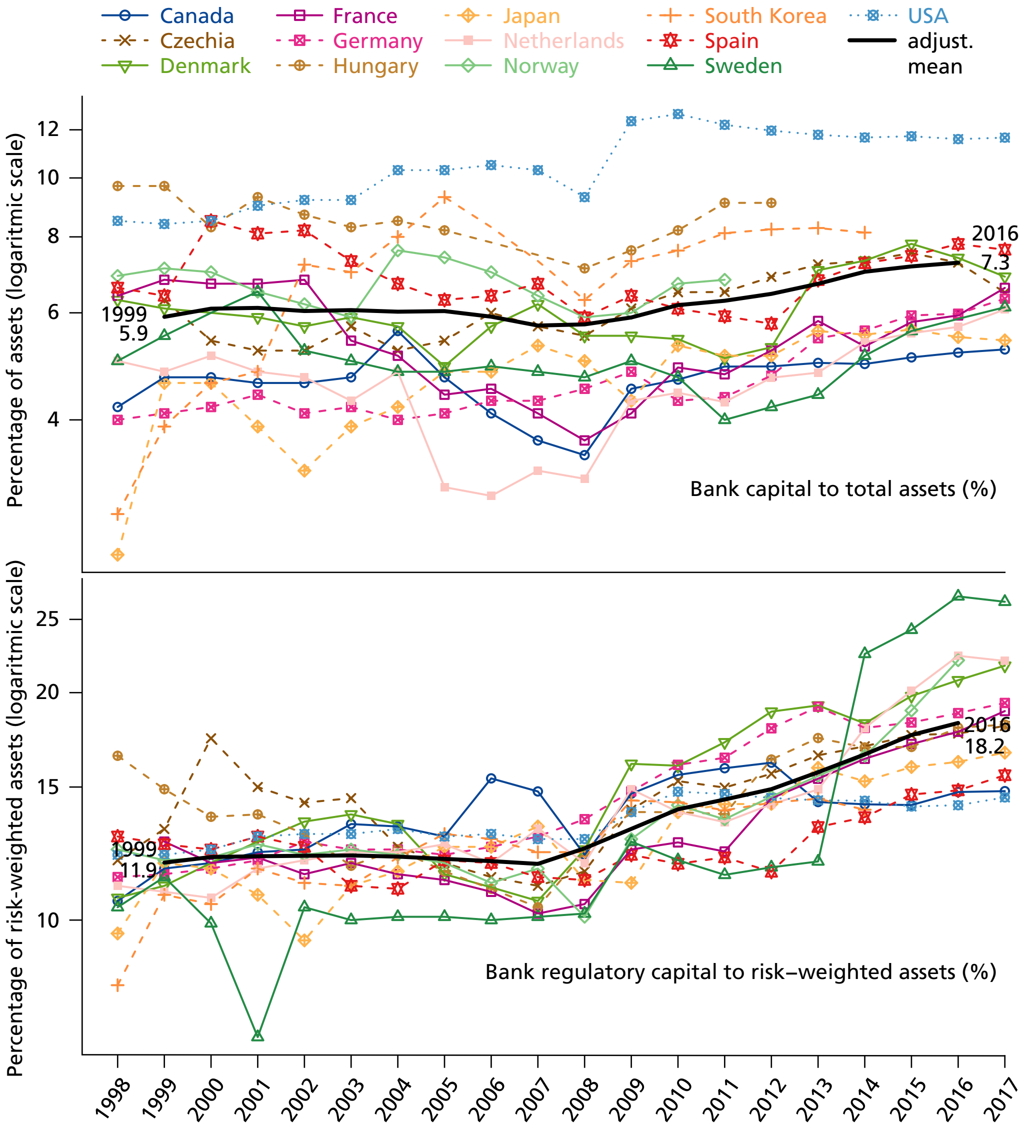
4 Finance, regulation, and inequality

Following the 2008 financial crisis, the approach to financial regulation around the world changed fundamentally. As early as 2009, both the Federal Reserve and the European Central Bank (ECB) expanded existing stress tests used internally into a general policy instrument. These stress tests model various economic crisis scenarios and mandate banks identified as risk-prone to raise more capital (Goldstein 2017; Tooze 2018). In addition, the crisis shock provided the impetus for the creation of the Basel 3 framework at the Bank for International Settlements, with national legislators transposing large parts of these recommendations into national law as of 2013 (Helleiner 2014). A particular focus in this accord was placed on increasing the quality and quantity of bank capital. This was done for instance with a minimum Common Equity Tier 1 ratio of 4.5 percent of risk-weighted assets and a minimum leverage ratio of Tier 1 capital to total assets of 3 percent without risk-weighting to reduce procyclical deleveraging.

The EU Commission introduced two-part financial legislation in 2013 consisting of the Capital Requirements Directive IV (CRD IV) and the Capital Requirements Regulation (CRR). The CRR stipulates, among other things, the Common Equity Tier 1 capital ratio of 4.5 percent of risk-weighted assets and the leverage ratio of 3 percent of Tier 1 capital, divided by total assets. The implementation of the Dodd-Frank Act in the United States has been a laborious and lengthy process but has led to the introduction of a capital ratio of 7 percent of risk-weighted assets and a leverage ratio of 5 percent, each of which goes beyond the European legislation to some extent (Acharya 2012; Eichengreen 2014; Tooze 2018).

The impact of regulations that tackle systemic risk and introduce higher capital requirements on earnings in the finance sector provides valuable insights into the state's ability to tackle earnings inequality through intervention in market mechanisms rather than post-production interventions such as income taxes. One might assume that higher capital ratios lead to lower bank profitability, which in turn leads to lower top earnings for financiers whose remuneration is linked to performance. In the following we test this hypothesis.

Figure 4 Bank capital ratios



In Figure 4, we depict the development of bank capital from 1998 to 2017. We can see an average increase in bank regulatory capital to risk weighted assets from 11.9 percent to 18.2 percent and a weaker increase in bank capital to total assets from 5.9 percent to 7.3 percent between 1998 and 2017. Before the crisis, however, on average bank capital stagnated or slightly decreased for most countries. In the few years before the crisis, in some countries – such as the United States, South Korea, Norway or Canada – there was a significant decrease in bank capital. Thus, we see that before 2008 on average the capital base of banks had eroded and, in some countries, substantially so.

After the crisis, the increase in bank capital is clearly visible. However, in most countries banks' key post-crisis capital adjustment was a restructuring of balance sheets toward assets that fall under the regulatory criteria, rather than increasing total capital. In some countries – for instance the United States, Canada, or France – there was a sharp increase for both indicators of bank capital after 2008, while most countries experienced a slight steady increase in bank capital over the post-crisis period. On average, however, we see a stronger post-crisis increase for regulatory capital to risk weighted assets than for bank capital to total assets. This could be due to a concomitant restructuring of banks' balance sheets toward less risk-prone assets and capital falling under the regulatory criteria. In Sweden, for instance, we see a significant increase after 2013 in regulatory capital to risk-weighted assets, but a relatively slow increase in capital to total assets, which may be due to a strong restructuring of Swedish banks' balance sheets. In most countries the immediate post-crisis adjustment in bank capital had a stronger effect on recapitalization than the introduction of higher capital requirements in the following years. For the United States, for instance, capital rose strongly in 2009 but even fell slightly over the subsequent years.

Capital requirements and inequality

To test the impact of the two capital ratios on both global inequality and finance's contribution to inequality, we apply the same methodology as in the previous section. We use equation 4, where we replace trading volume by one of the two capital ratios (Table 3). We further introduce an interaction term of these variables with the period (2009–2017) during which banks strengthened their capital base.

Our first ratio – bank capital to total assets – does not have any significant impact on inequality (Model 3.1 and 3.2), but it does have a small significant impact on finance's contribution to inequality (Model 3.1). One standard deviation increase in the bank capital to assets ratio decreases finance's contribution to inequality by 0.10. When we decompose the period into pre- and post-crisis, we see that the stronger capital requirements have a slightly stronger impact than during the pre-crisis period, with an additional –0.04 effect. However, this marginal effect is not significant. Hence, the modest increase in capital ratios, as shown in Figure 4, had a modest effect in reducing the contribution of finance to inequality and no effect on reducing overall inequality. This mitigating effect remains of a similar magnitude to the increase in leverage during the upswing period.

Conversely, when we use the *bank regulatory capital to risk-weighted assets* ratio, we do not find any impact on our two measures of inequality. This contrast suggests that increasing the sophisticated *bank regulatory capital to risk-weighted assets* is less effective than increasing the crude *capital to assets* ratio. Of course, regulatory capital requirements targets were implemented to reduce risk-taking rather than top financiers' remuneration. Yet top financiers' remuneration may be a proxy of banks' risk-taking,

Table 3 Impact of capital ratios on inequality

	Top 1% share		Finance earnings' over-representation in top 1%	
	(1)	(2)	(3)	(4)
GDP per capita	0.05	0.05	0.49***	0.49***
	(0.06)	(0.06)	(0.13)	(0.13)
Union rate	0.05	0.05	0.21**	0.21**
	(0.04)	(0.04)	(0.08)	(0.09)
Importation rate	0.05	0.05	–0.16	–0.16
	(0.07)	(0.07)	(0.13)	(0.13)
Bank capital to total assets	0.04	0.03	–0.10*	–0.09
	(0.05)	(0.06)	(0.05)	(0.06)
Bank capital to total assets × (Years > 2008)		0.03		–0.04
		(0.10)		(0.14)
Country fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
R^2 (full model)	0.58	0.58	0.52	0.52
R^2 (proj model)	0.01	0.01	0.26	0.26
Num. obs.	185	185	167	167
Num. groups: country	13	13	12	12
Panel B.				
Bank regulatory capital to risk-weighted assets	–0.02	0.02	0.10	0.04
	(0.07)	(0.09)	(0.08)	(0.10)
Bank regulatory capital to risk-weighted assets × (Years > 2008)		–0.08		0.13
		(0.10)		(0.11)
Control variables	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
R^2 (full model)	0.57	0.57	0.54	0.54
R^2 (proj model)	0.00	0.01	0.31	0.32
Num. obs.	189	189	171	171
Num. groups: country	13	13	12	12

Note: OLS models with country and year fixed effects and panel corrected standard errors in parenthesis. Independent variables are one-year lagged. Dependent and independent variables are country-demeaned standardized. Hence, one standard deviation of bank capital ratio decreases by 0.10 standard deviation finance earnings' overrepresentation in the top 1% earnings share. ***$p<0.01$; **$p<0.05$; *$p<0.1$

and a sophisticated, tailor-made measure of capital used in banking activity is open both to inventive bookkeeping and to a politicization of accounting. Thus, our evidence suggests that risk does not decrease significantly when the sophisticated ratio of banks' regulatory capital to risk-weighted assets increases.

The bonus cap

In 2010, the European Commission proposed a first crisis-inspired directive, also known as the Capital Requirements Directive III (CRD III), which introduced a first bonus regulation. This bonus regulation set a minimum of 40 to 60 percent of variable remuneration that should be deferred over three years and a minimum of 50 percent that should be paid in instruments such as shares.[10] The aim of this directive was to curb excessive risk-taking by bankers, which the Commission attributed to the high bonuses paid in the case of good performance. With CRD IV/CRR, the Commission introduced a bonus cap based on the first bonus regulation, which went significantly beyond that, however. Implemented as a directive with CRD IV, the bonus cap set a maximum variable to fixed remuneration ratio of 100 percent and 200 percent if the shareholders approve. Additionally, the Commission, together with the EBA, introduced stricter guidelines for banks to identify regulated staff in 2014 to make implementation more uniform among member states.

Although a bonus regulation was introduced with CRD III, it is with the fixed/variable remuneration ratio set by CRD IV in 2013 that the bonus cap promises to be truly effective for reducing variable remuneration. The probable adjustment of banks' remuneration practices to the bonus cap is to reduce variable remuneration in the short term to comply with the requirements and increase their employees' fixed remuneration (Murphy 2013). The effect on total remuneration is less unequivocal, as it could adhere to two alternative hypotheses. On one hand, risk-averse employees might accept a reduction in total remuneration in exchange for a less risky composition of their remuneration. On the other hand, financiers' strong bargaining power (Guillot 2021) could prevent any attempt to reduce earnings and leave absolute remuneration levels unchanged.

We will thus examine these hypotheses in this section. One difficulty that arises in the analysis is that, as noted above, the requirements as to whose remuneration must be disclosed in bank reports were made more stringent in 2014. Consequently, the number of material risk takers whose remuneration is disclosed in annual bank reports almost doubled (exp (0.64)) (Table A4). This indicates that we need to take the larger perimeter of the material risk takers into account, to circumvent a downward bias of the impact of the bonus cap on earnings. The expansion of the population through the more uniform measures to identify material risk takers leads to the inclusion of lower ranking financiers, who earn less. To avoid such bias, we control for the impact of the change in size and check for the robustness of our estimates in a variety of functional forms.

In the following, we use OLS regressions to analyze the effect of the bonus cap on the remuneration practices of European banks. In the models, we include three dependent variables: the log of fixed, variable, and total remuneration per material risk taker in each "job."[11] A "job" is defined by the intersection of a bank and a particular function of employees within that bank. As we use aggregate statistics per job, we weight observations with the number of material risk takers they represent. This prevents giving

10 The directive applied to all credit institutions as defined in Directive 2006/48/EC and investment institutions as defined in Directive 2006/49/EC. With this the Commission largely followed the recommendations of the Financial Stability Forum and from the start of 2011 the CRD III was to be implemented by member states.

11 All values of independent and dependent variables are in 2015 prices and converted to euros based on ECB exchange rates.

Table 4 Bank fixed effects for logged remuneration

	Model 1			Model 2		
	Fixed	Variable	Total	Fixed	Variable	Total
Regulated period	0.18**	–0.85	–0.02			
	(0.08)	(0.53)	(0.04)			
Total assets (log)	0.11	1.01***	0.38***	0.40**	2.81***	0.84***
	(0.09)	(0.33)	(0.10)	(0.17)	(0.67)	(0.17)
Earnings before interest and taxes (asinh)	0.00	0.02	0.01***	0.00	0.01	0.00
	(0.00)	(0.02)	(0.00)	(0.00)	(0.02)	(0.00)
Employees (log)	0.02**	0.30*	0.04***	0.01	0.17**	0.02
	(0.01)	(0.16)	(0.01)	(0.01)	(0.07)	(0.01)
Material risk takers	–0.36***	–0.97***	–0.57***	–0.32***	–0.52**	–0.53***
	(0.09)	(0.20)	(0.03)	(0.10)	(0.24)	(0.05)
Material risk takers (squared)	0.03***	0.12***	0.05***	0.03**	0.07**	0.05***
	(0.01)	(0.02)	(0.00)	(0.01)	(0.03)	(0.01)
2009				–0.14	–3.11***	–0.27***
				(0.10)	(0.26)	(0.06)
2010				–0.11*	–1.17**	–0.01
				(0.05)	(0.45)	(0.07)
2011				–0.12**	–1.54***	–0.22***
				(0.05)	(0.31)	(0.06)
2012				0.02	–0.31	0.01
				(0.04)	(0.20)	(0.05)
2014				0.14*	–1.30***	–0.06
				(0.07)	(0.39)	(0.04)
2015				0.24***	–0.96***	0.07
				(0.06)	(0.34)	(0.05)
2016				0.27***	–0.94***	0.05
				(0.06)	(0.28)	(0.09)
2017				0.27***	–0.40*	0.19***
				(0.06)	(0.23)	(0.06)
R^2 (full model)	0.96	0.92	0.97	0.96	0.95	0.98
R^2 (proj model)	0.65	0.42	0.66	0.68	0.66	0.73
Num. obs.	405	364	428	405	364	428
Num. groups: jobs	69	66	72	69	66	72

Note: OLS models with job fixed effects, weighted by number of material risk takers and clustered robust standard errors in parentheses. Number of material risk takers in thousands. ***p<0.01; **p<0.05; *p<0.1

disproportionate weight to observations with a small number of regulated employees. In all models, we introduce "job" fixed effects. We thus control for time invariant heterogeneity bias, in other words, for bank and job-specific differences that we cannot easily quantify, such as in corporate governance standards, human resource practices, shareholder fragmentation or the like.

In Model 1, Table 4, we use the regulation period (2014–2017) as our main independent variable. It captures the average within-job change in financier's remuneration during the regulated period (from 2014 on) in comparison with the unregulated period (before 2013). In model 2, year-dummies are taken as the main independent variables, with 2013 as the reference category. With this we measure the yearly changes in remuneration before and after the introduction of the bonus cap. The second model thus addresses the short-run change in remuneration practices induced by the bonus cap. The analysis includes controls for total assets and number of total employees to control for bank size, which could have a positive effect on executives' remuneration. We further add earnings before interest and taxes to control for banks' profitability, which could have a positive effect on executives' remuneration. Finally, we introduce a quadratic control specification of the number of material risk takers to account for the potential increase in material risk takers after 2013. In Table A5, we have tested whether our results are robust when we change our specification.

As predicted, the bonus cap led to a substantial increase in fixed wages. Model 1 in Table 4 shows that there is a 20 percent increase in fixed wages (that is, exp (0.18)) over the 2014–2017 regulatory period. Model 2 further specifies that fixed wages initially increased by 15 percent in 2014 and by an additional 12 percent in 2015 and did not decrease thereafter.

Conversely, the impact of the bonus cap on variable remuneration is less precise. Model 1 estimates a strong −57 percent negative effect (that is, exp (−0.85)−1) which is insignificant at conventional level (p = 11 percent). In addition, Table A5 shows that the estimate of this parameter varies by specification and is not significant despite a high absolute value. Model 2 provides a more accurate picture of the annual evolution of bonuses and facilitates an understanding of the results of Model 1. In 2014, bonuses dropped substantially (−73 percent, that is, exp (−1.30)−1) and significantly compared with their 2013 level. This short-term effect was attenuated in subsequent years, with log parameters falling from −1.3 to −0.96 and −0.94. That the parameter for variable remuneration and regulated period in Model 1 is not significant is also due to the evolution of bonuses before 2012. Bonuses were significantly lower in 2011 than in 2014. Therefore, the evaluation of the effect of the bonus cap depends on the determination of the reference period to which we compare the regulated period. If we take only the years 2013 or both 2012 and 2013, there is a significant decrease in bonuses with the introduction of the bonus cap. If we take the period 2009–2013 as the reference period, no clear decline can be seen. In the latter case, the inclusion of years of deep crises with the 2009 GFC and 2010–2011 sovereign debt crisis may bias the reference period. Compared with the more normal years, it is therefore likely that there was a decrease in bonuses due to the bonus cap.

Finally, we show the impact of the bonus cap on total remuneration. The results from our models show that the bonus cap did not lead to any significant change in total remuneration. This "null" result holds true for almost all specifications (Tables 4, A5).

Hence, the introduction of the bonus cap did lead to a robust increase in fixed remuneration of approximately 20 percent. The effect on variable remuneration is more difficult to interpret. While we measure a strong short-term drop in variable remuneration in 2014, this decline is far from robust. This inconsistency could be due to the difficulty of simultaneously accounting for a change in remuneration and a change in the perimeter of material risk takers. It could also be because we do not capture the business climate determinants of financiers' remuneration. An alternative explanation is that banks may, for a short time, have complied with the new rules by reducing bonuses or using inventive accounting techniques, loosening their compliance later on. In the United

Kingdom, for instance, banks have turned bonuses into so-called "role-based allowances," which they booked as fixed remuneration, and which thus do not appear as variable remuneration in their reports but continue to serve this purpose.[12] Finally, although the change in earning composition increases banks' salary costs because fixed salaries are less downward flexible and should have led banks to reduce overall remuneration, banks have not adjusted in this direction.

In sum, apart from the modest restructuring of European banks' remuneration practices by reducing variable remuneration and increasing fixed remuneration, a real reduction in top earnings in banks cannot be observed. Banks adapted to the new rules and aligned previous total remuneration packages by increasing fixed remuneration. Of course, we cannot exclude the impact of other time-variant, confounding variables that capture the business climate. Nevertheless, it seems plausible that these developments are a result of the bonus cap because other plausible explanatory mechanisms are insignificant in our analysis. Therefore, we can assume with high confidence that the bonus cap had no impact on reducing top earnings and thus earnings inequality.

5 Conclusion

Our study sheds new light on the finance–inequality nexus, with four distinct contributions. First, we corroborate with detailed administrative data for a large set of high wage countries that the upswing in finance was a major driver of the increase in earnings inequality. This process occurs through a simple and powerful mechanism, the granting of high wages in a growing financial industry. Second, we show that the significant downswing in finance after the crisis did not contribute to a symmetrical decline in earnings inequality. Third, we have shown that banks have significantly increased their capital since 2008 and this may well be influenced by higher capital requirements. However, this increase in bank capital did not contribute to a significant decline in top earnings and fostered only a modest decline in the overrepresentation of financiers among the top 1 percent earnings share. Fourth, we have shown that the introduction of the bonus cap in the EU contributed to a short-term decrease in variable remuneration and a consistent increase in fixed remuneration, while the total remuneration of top financiers has remained unaffected by the bonus cap.

Hence, neither the post-financial crisis downswing in financial activity nor financial regulation have contributed to a significant reduction in inequality to the extent that the 1990–2007 financial upswing contributed to its increase. The contribution of finance to inequality in times of upswing thus has long-term and hardly reversible effects.

12 Cf. Schäfer and Arnold (2014) and EBA (2015).

Further research needs to explore the roots of this asymmetry in the evolution of financiers' earnings in more detail. First, it is worth analyzing the shift in the dominance of traditional banks in the financial sector to non-bank financial institutions operating in unlisted markets, as well as to private equity and shadow banks. These non-traditional forms of finance are central to financialized capitalism, as noted in the recent financialization literature. If this trend proves predominant, wage resilience in finance may also stem from alternative engines of value creation.

Qualitative studies of banks' remuneration practices provide valuable insights into possible mechanisms of financial wage resilience (Godechot 2017). When team leaders evaluate the performance of their supervisees for the distribution of bonuses, they assign asymmetric responsibility for gains and losses. Profit, which is actively sought, is considered a responsibility and an achievement of the financiers, while losses, which are not sought, are considered a matter of bad luck. It is thus common that considerable bonuses are paid despite high trading losses, especially if the responsible financiers are promising talents and their sectors of activity are booming in the market. Moreover, when repeated dismal performance in a business unit forces cost-cutting, banks tend to cut the bonuses of younger bank employees and maintain the salaries of their star performers, whom they deem essential. This asymmetry could be exacerbated by the strong bargaining power of finance employees, who can shift their activities to a competitor by taking technology, customers, colleagues, and subordinates with them. This may be further promoted if employees indeed give priority to the fight against wage cuts over the fight for wage increases (Simiand 1931; Keynes 1936; Kahneman, Knetsch, and Thaler 1986). Downward wage rigidity and stickiness in finance could therefore be a long-term driver of global inequality.

Finally, this research raises policy issues. By highlighting the asymmetry in the contribution of finance to earnings inequality in times of financial upswings and downswings, we show that the structural redistribution of earnings through financialization is not readily reversible. The fact that finance has lost prominence in public discourse over the past decade and has been overshadowed by the incomes and wealth of tech superstars does not mean that it no longer plays a significant role in earnings inequality. Policymakers seeking to curb earnings inequality must therefore be mindful of the special role played by the financial industry. Policymakers should consider more effective measures to reduce excessive wages than the capital requirements and bonus cap analyzed in this paper.

Bank capital regulations should avoid overly sophisticated risk-weighted capital measures, which banks can use creatively to circumvent the intention behind the policy. Crude measures for capital ratios, although less sophisticated, might be more suitable. Bonus caps should be complemented by regulation of total wages, either through a cap or a specific and permanent payroll tax for top earners in the financial sector.[13]

13 One of the main shortcomings of the "75 percent tax" for millionaires in France introduced by the socialist Hollande government is that it was announced from the beginning that this tax would be introduced for only two years (Guillot 2021). Therefore, most firms avoided reforming their remuneration scale and simply paid the bulk of the tax.

Furthermore, these measures should not be limited to traditional banks but extended to all financial institutions. Finally, preventing an upswing in the financial sector as a whole or in a submarket could be a way to avoid a permanent distortion of the earnings scale.

Appendices

A1 Data description

Canada (1990–2013). Data were generated by Statistics Canada. The data are population-level and include all sectors and industries and employees.

Czechia (2002–2016). Data were taken from the Average Earnings Information System (ISPV) survey conducted by the private agency TREXIMA. The data consist of the entire population of public sector workplaces, plus a sample of private sector workplaces. The private sector sample consists of workplaces with at least 10 employees. A stratified sampling of private sector workplaces with 10–250 employees was taken based on the size of the workplace. All private sector workplaces with over 250 employees are included in the data. The data also spans all industries and sectors. In the end, the dataset covers 80 percent of the Czech workforce and 96 percent of the workforce in establishments with 10 or more employees. Estimates are weighted to correspond to the complete workforce in establishments with 10 or more employees.

Denmark (1994–2015). The data consist of population-level observations of both private and public sector workplaces extracted from the labor market statistics register (Den Registerbaserede Arbejdsmarkedsstatistik – RAS), and earnings from the job register IDAN. Demographics such as age, gender, and nativity come from the population register (Befolkningsregistret).

In order to exclude marginal jobs, we exclude workers earning less than half the P10 threshold of full-time workers. The wage data show a discontinuity in 2008: the introduction of the e-income register in 2008 substantially increased the number of marginal jobs in the register. In order to consistently select the working population, we matched the 2008 threshold on 2007 by applying a +18,000 Kr correction in 2008 and following years.

France (1993–2019). Our analyses use data from the DADS social security register *(Déclaration annuelle de données sociales)*. Access to the DADS data was obtained through the CASD *(Centre d'accès sécurisé aux données)*, for researchers authorized by the French *Comité du secret statistique*. The data consist of population-level observations of private sector workers. State civil servants are missing before 2009 and excluded in the following years for the sake of consistency.

Germany (1990–2015). Data come from a customized sample for the project "Dynamics of Organizational Earnings Inequality: Investigation within the Comparative Organizational Inequality International Network (COIN)" of the Integrated Employment Biographies Sample (IEBS) of the Federal Employment Agency. It covers roughly 5 percent of the German working population and about 20,000 establishments, spanning the years 1999–2015. Estimates are weighted to correspond to the complete workforce.

Earnings not subject to social security contributions because they are below the threshold for small-scale employment (for example, newspaper delivery) – which is currently 450 euros per month – are excluded from the sample. The earnings are also top-coded at the social contribution limit, which differs by year and for eastern and western Germany. To impute the top-coded earnings, an imputation strategy based on the imputation from Card, Heining, and Kline (2013) was established, which accounts for individual and establishment wages prior to the censored period. However, rather than focusing on the mean individual and establishment wage prior to the censored observation, as was done by Card, Heining, and Kline, we utilize information on lagged earnings. Given the limitation of our imputation, measures of exposure involving the top 1 percent should therefore be considered with caution.

Hungary (2003–2011). Our analyses use Admin2 data processed by the Institute of Economics, Centre for Economics and Regional Studies of the Hungarian Academy of Sciences. These data are generated by linking data from five governmental institutions (the Pension Directorate, the Tax Office, the Health Insurance Fund, the Office of Education, and the Public Employment Service). The data are a 50 percent random sample of the Hungarian population from 2003 to 2011. The earnings concept is monthly earnings from each person's primary job. Monthly data were aggregated to obtain yearly wages. Low-wage workers, defined as workers earning less than half of the yearly minimum wage, are dropped from the sample.

Japan (1989–2013). Data are from the Basic Survey on Wage Structure conducted by the Ministry of Health, Labor, and Welfare of Japan. The survey is a two-stage design in which a sample of private sector establishments with at least five employees are selected, and then a uniform random sampling of workers among these establishments is taken. Firms' executives are not included in the data. Given this limitation and the small size of the sample, measures of exposure involving the top 1 percent should therefore be considered with caution, but 10 percent thresholds are treated as more reliable. The sample covers 4 percent of the workforce working in establishments with more than five workers. Estimates are weighted to correspond to the complete workforce.

The Netherlands (2006–2018). Annual data on employee wages and company sector and industry are provided by Statistics Netherlands (CBS) within the System of Social-Statistics Database (SSB). We linked data on employees and employing firms to construct a dataset with population-level coverage of wages across all sectors and industries. The analyses include the highest-paying jobs of each employee in a given year, and jobs with wages lower than age-specific minimal hourly wages are excluded.

Norway (1996–2014). Data were generated by Statistics Norway and are population-level, including all sectors and industries, although private sector identifiers are available only beginning in 1999.

South Korea (1982–2012). Data are from a survey conducted by the Korean Ministry of Labor. The data consist of a sample of private sector establishments, first stratified by size and then by region and industry. An establishment must have had a minimum of five employees to be included in the sample before 1999, and ten employees beginning in 1999. All industries except agriculture are included. The dataset contains only full-time jobs. Estimates are weighted to produce national estimates.

Spain (2006–2016). Our analyses use data from the Continuous Sample of Working Histories (CSWH; *Muestra Continua de Vidas Laborales con datos fiscales*) from Spain's Social Security Office. The CSWH contains matched anonymized social security, income tax and census records for a 4 percent non-stratified random sample of the population who in one specific year had any connection with Spain's social security system (whether via employment, self-employment, unemployment, or retirement). The CSWH provides information on individuals' complete labor market histories from 1980 (or the year the individual registered with Social Security) to the year of data collection.

Because earnings from the social security records are top- and bottom-capped, we use earnings from tax records containing uncensored gross labor earnings for each job (tax records are available from 2006 onwards). Thus, the procedure is as follows: first, we identify personal information from social security records, then match those records with the individuals in the tax dataset, thereby obtaining 2006–2017 earnings from tax records. Consequently, we use the full information on the labor market history of individuals to compute their tenure and other variables but study earnings only for the years 2006 to 2017, for which tax data are available. When multiple jobs overlap, we only consider the main job, which is either that with the longest spell within the same firm or that with the highest earnings across firms. In this way, we build a yearly panel that covers job spells, with a start/end date and tied to a firm identifier.

Sweden (1990–2012)**.** The data used are from population-wide administrative registers from Statistics Sweden (the LISA database) and cover all sectors and industries. However, occupations are available only after 2001 and hourly wages are not available.

A2 Supplementary tables and figures

Table A1 Presentation of linked-employer administrative data

Country	Start year	End year	Field	Definition of threshold	Threshold wage (end year)	Number workers (end year)	Source
Canada	1990	2013	Exhaustive	1/2 full-time full year minimum wage	7857 CAD	18,943,396	Statistics Canada
Denmark	1994	2015	Exhaustive	1/2 full-time P20	48,178 DKK	2,261,464	RAS, IDAN and BES
Norway	1996	2014	Exhaustive	1/2 full-time P10	60,957 NOK	2,038,835	Statistics Norway
Sweden	1990	2012	Exhaustive	1/3 prime age P50	93,210 SEK	4,192,204	Statistics Sweden
France	1993	2019	Exhaustive private sector	1/2 full-time full year minimum wage	7,688 EUR	19,263,900	DADS
Germany	1990	2015	Sample of workers (6 %) in 20,000 establishments	1/2 full-time P10	12,871 EUR	1,120,354	IEBS
Netherlands	2006	2018	Exhaustive	Age-specific minimum hourly wage	4 EUR per hour	8,867,793	CBS
Spain	2006	2016	Random sample of workers born since 1962 (4 %)	1/2 full-time full year minimum wage	2,799 EUR	380,804	Continuous Sample of Working Histories (CSWH) and tax records
Czechia	2002	2016	Sample of workers (80 %)	1/2 full-time full year minimum wage	52,830 CZK	1,917,812	Average Earnings Information System (ISPV) survey
Hungary	2003	2011	Sample of workers (50 %)	1/2 full-time yearly minimum wage	468,000 HUF	1,017,665	Admin2
South Korea	1990	2012	Sample of workers (8 %) out of a sample of private sector establishments size>5	1/2 full-time full year minimum wage	4,763,200 KRW	613,369	Korean Ministry of Labor
Japan	1989	2013	Sample of workers (4 %) out of a sample of private sector establishments of size>5	1/2 full-time P10	1,056,700 Yen	1,089,517	Basic Survey of Wage Structure
USA	1990	2016	Exhaustive	No threshold	0 USD	148,658,000	SSA in Piketty Saez (2003) updated (2020)

Table A2 Banks in sample

Bank	Country	Start year	End year
ABN AMRO group	Netherlands	2011	2017
AIB group PLC	Ireland	2011	2017
Banco Santander SA	Spain	2011	2017
Bank of Ireland group PLC	Ireland	2010	2017
Bankia SA	Spain	2012	2017
Barclays PLC	United Kingdom	2013	2017
BBVA	Spain	2011	2017
BNP Paribas	France	2009	2017
Commerzbank	Germany	2010	2017
Danske Bank AS	Denmark	2011	2017
Deutsche Bank AG	Germany	2010	2017
Erste Group Bank AG	Austria	2012	2017
HSBC HLDGS PLC	United Kingdom	2013	2017
ING Groep NV	Netherlands	2012	2017
Intesa Sanpaolo SPA	Italy	2011	2017
Jyske Bank	Denmark	2011	2017
National Bank of Greece	Greece	2013	2017
NATIXIS	France	2010	2017
Permanent TSB Group HLDGS	Ireland	2014	2017
Skandinaviska Enskilda Bank	Sweden	2010	2017
Société Générale Group	France	2010	2017
Svenska Handelsbanken	Sweden	2010	2017
Swedbank AB	Sweden	2009	2017
Sydbank AS	Denmark	2011	2017
Unicredit SPA	Italy	2009	2017

Table A3 The asymmetric effect of capitalization on inequality and finance's contribution to inequality

	Top 1% share			Finance earnings' overrepresentation in top 1%		
	(1)	(2)	(3)	(4)	(5)	(6)
GDP per capita	0.06	0.08*	0.12**	0.59***	0.61***	0.62***
	(0.04)	(0.04)	(0.05)	(0.10)	(0.10)	(0.10)
Union rate	0.18***	0.16***	0.21***	0.09	0.08	0.10*
	(0.04)	(0.04)	(0.05)	(0.06)	(0.06)	(0.05)
Importation rate	–0.08	–0.04	0.08	–0.09	–0.07	–0.03
	(0.07)	(0.07)	(0.06)	(0.08)	(0.08)	(0.08)
Capitalization/GDP	0.31***	0.41***	0.49***	0.61***	0.73***	0.67***
	(0.08)	(0.08)	(0.10)	(0.14)	(0.20)	(0.15)
Capitalization/GDP × Years in (2001–2003, 2008–2013)		–0.34**			–0.37*	
		(0.14)			(0.21)	
Cumulative sum of drops in capitalization/GDP			–0.42***			–0.15*
			(0.11)			(0.08)
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
R^2 (full model)	0.64	0.65	0.69	0.54	0.55	0.54
R^2 (proj model)	0.15	0.17	0.25	0.37	0.38	0.38
Num. obs.	254	254	254	227	227	227
Num. groups: country	13	13	13	12	12	12

Note: OLS models with country and year fixed effects and panel corrected standard errors in parenthesis. Dependent and independent variables are country-demeaned standardized. All independent variables are one-year lagged. ***p<0.01; **p<0.05; *p<0.1

Table A4 Job and bank fixed effects for logged number of material risk takers

	Log of the number of material risk takers declared	
	Job FE	Bank FE
2009	–0.32	–0.08
	(0.42)	(0.50)
2010	0.11	0.21
	(0.22)	(0.39)
2011	0.23	0.14
	(0.14)	(0.23)
2012	0.08	0.10
	(0.09)	(0.16)
2014	0.64***	0.62***
	(0.10)	(0.13)
2015	0.62***	0.77***
	(0.12)	(0.13)
2016	0.59***	0.72***
	(0.13)	(0.17)
2017	0.55***	0.56***
	(0.15)	(0.19)
Total assets (log)	–0.83**	0.31
	(0.39)	(0.52)
Earnings before interest and taxes (asinh)	–0.02***	–0.01
	(0.01)	(0.02)
Number of employees in the bank (log)	–0.06***	–0.02
	(0.02)	(0.03)
Position fixed effects	Yes	No
Bank fixed effects	No	Yes
R^2 (full model)	0.92	0.81
R^2 (proj model)	0.32	0.18
Num. obs.	428	171
Num. groups: Jobs	72	
Num. groups: Banks		25

Note: OLS models with job or bank fixed effects and cluster robust standard errors in parenthesis. ***p<0.01; **p<0.05; *p<0.1. 2013 is used as reference category.

Table A5 Robustness checks: Impact of regulation on pay according to various specifications

Specif.	Control variables	Fixed	Variable	Total
	A. Regulated period parameter			
0	None	0.30*** (0.08)	–0.57* (0.31)	0.14 (0.11)
1	Total assets, earnings, size of the banks	0.29*** (0.10)	–0.43 (0.48)	0.17 (0.11)
2	Idem+Nb risk takers	0.28*** (0.06)	–0.43 (0.46)	0.16** (0.07)
3	Idem+(Nb risk takers)^2	0.18** (0.08)	–0.85 (0.53)	–0.02 (0.04)
4	Idem+(Nb risk takers)^3	0.21** (0.09)	–0.52 (0.47)	0.01 (0.05)
5	Idem+(Nb risk takers)^4	0.20** (0.09)	–0.69 (0.51)	0.02 (0.05)
	B. 2014 dummy variable parameter			
0	None	0.13** (0.06)	–1.27*** (0.35)	–0.08 (0.08)
1	Total assets, Earnings, size of the banks	0.14** (0.06)	–1.27*** (0.32)	–0.05 (0.09)
2	Idem+Nb risk takers	0.15** (0.06)	–1.27*** (0.32)	–0.05 (0.08)
3	Idem+(Nb risk takers)^2	0.14* (0.07)	–1.30*** (0.39)	–0.06 (0.04)
4	Idem+(Nb risk takers)^3	0.17** (0.08)	–1.06*** (0.37)	–0.05 (0.04)
5	Idem+(Nb risk takers)^4	0.17** (0.08)	–1.06*** (0.37)	–0.05 (0.03)

Note: OLS models with job effects and cluster robust standard errors in parenthesis.
***p<0.01; **p<0.05; *p<0.1

Figure A1 Banks' return on equity before tax

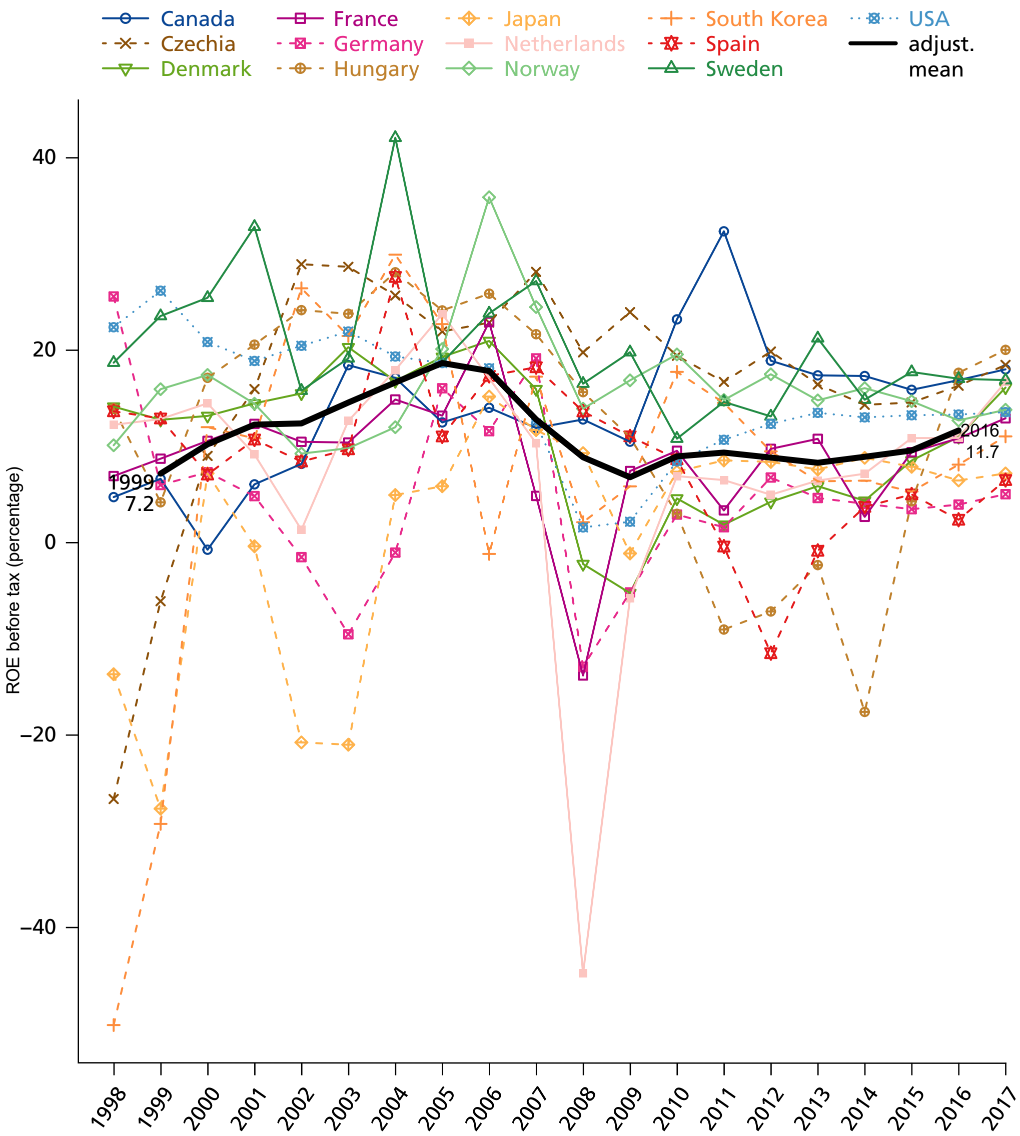
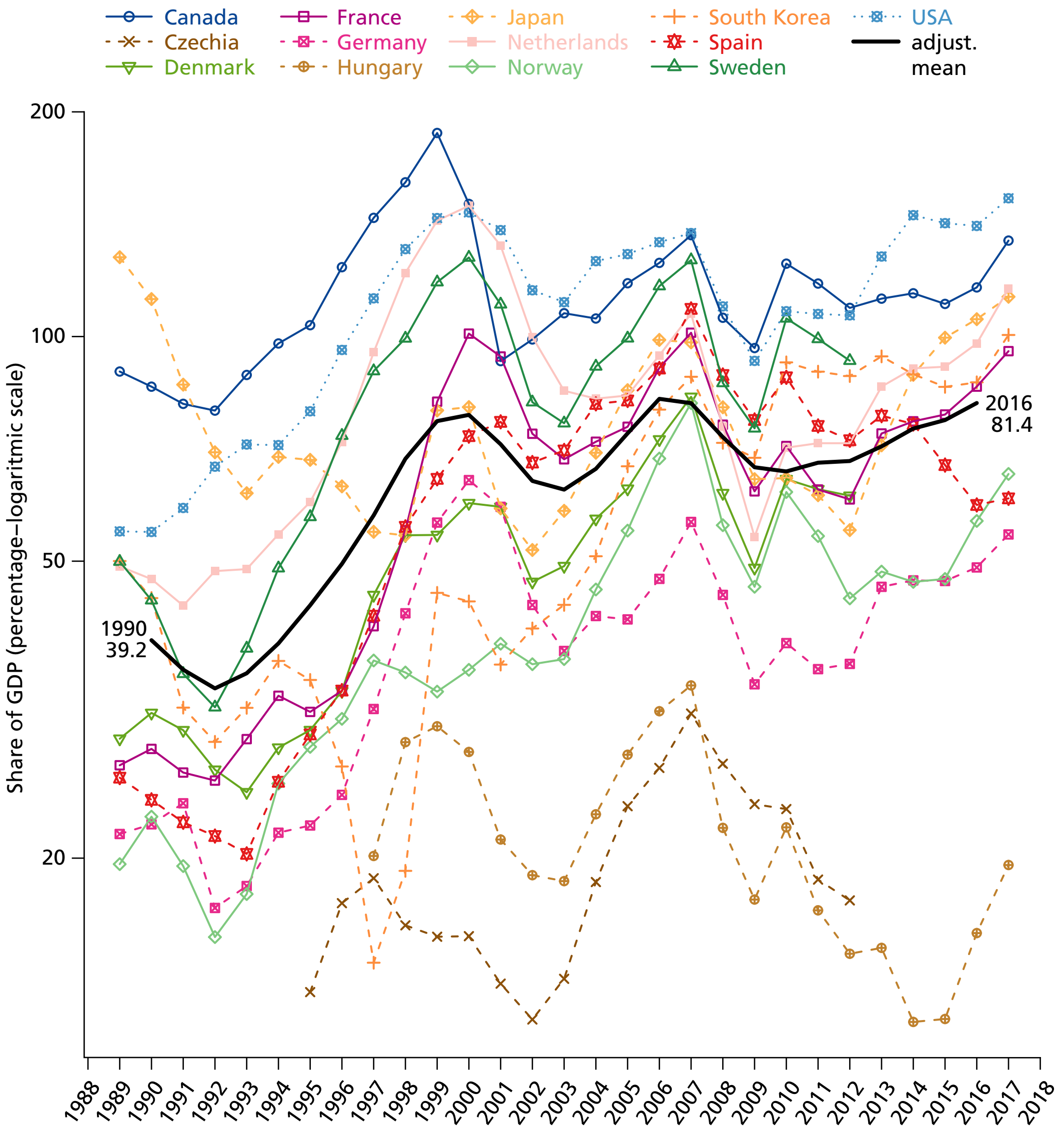


Figure A2 Capitalization to GDP



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