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FINANCIAL SECTOR TRANSFORMATION AND INCOME INEQUALITY– AN EMPIRICAL ANALYSIS

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Abstract

This paper examines the relationship between financial sector transformation and income inequality. We construct an econometric model of income concentration for a panel of 16 OECD countries in the years 1995-2009. From our study, financial sector transformation, measured individually by three indicators (GDP share of stock market value traded, bank income and private credit), emerges as a nexus of complex and interconnected phenomena, which are strongly associated with the concentration of income at the top of the distribution.

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Introduction

This paper examines social consequences of the rising importance of financial sectors in the real economy. It analyses the distributive forces generated by financial sector transformation. Understood as changes in financial intermediation (i.e. channelling of funds between lenders and borrowers by bank and non-bank intermediaries in a financial system), financial sector transformation is often described as "financialisation". Epstein (2005) defines financialisation as the "increasing role of financial motives, financial markets, financial actors and financial institutions in the operation of the domestic and international economies" (p. 3). It is an extremely complex process occurring within a variety of dimensions. Although most pronounced in the USA, financial sector transformation has also taken place in various aspects and at different points since the 1980s in Europe (cf. Veronese Pasarella, 2013).

Financialisation finds its roots in the persistently high inflation and interest rates in the late 1960s, which induced non-financial firms to seek investment financing through financial markets rather than banks. This realigned firms' objectives away from long-term investment towards short-term profitability, making them more involved in financial activities (such as issuing shares). This raised the importance of financial over real profits (Palley, 2007, p. 18). Such changes in corporate behaviour contributed to the growing share of the financial sector in the economy at the expense of manufacturing.

Financial sector transformation gained steam in the 1980s under policies promoting market liberalisation and retrenchment of the state from public service provision associated with the government of Reagan in the USA and Thatcher in the UK (Sawyer, 2013, p. 13). Liberalisation of labour markets and the resultingscaling back of minimum wages, unemployment protection schemes and union-oriented policies resulted in a gradual decline of wage income growth. Simultaneously, provision of pensions, housing and public goods such as education and healthcare became delegated to the private sector. With stagnant wages and diminishing state provision, households found themselves in need of additional financing through borrowing.

Rising credit demand was paralleled by the massive proliferation of financial instruments and the development of structured finance. The turn of non-financial companies towards financial markets resulting from high borrowing costs in the 1960s and 70s led financial intermediaries to seek revenue in the household sector and through innovation of new financial products (Dymski, 2009, p. 157). An increasing volume of financial obligations — primarily mortgages and consumer debt – was transformed into securities in a process of securitisation, forming collateralised debt obligations (CDOs), which combined financial instruments of varying risk and return characteristics (Pollin & Heintz, 2013, p. 113). The establishment of credit default swaps (CDS) and derivatives on existing products allowed investors to bet against the default of any financial instrument, leading to the transformation of traditional lending relations based on intermediation towards an "originate and redistribute" model, where default risk became "originated" by creditors and then spread across the financial system through securitisation. This new lending model was adopted by not only registered banks, transformed into highly consolidated "megabanks" as a result of intense merger activity, but also non-bank intermediaries, which played a role similar to that of formal banks but were outside the central bank's jurisdiction in obtaining liquidity (Pollin & Heintz, 2013, p. 115). This process was validated by increasing financial deregulation policies, which allowed commercial banks to engage in financial investment activities.

This paper argues that these processes associated with financial sector transformation exerted a direct impact on income distribution in advanced countries. There has recently been an upsurge in studies documenting the dramatic rise of income inequality around the world (cf. Piketty, 2014; Alvaredo, Atkinson, Piketty & Saez, 2013). In the USA, where the trends are the most extreme, the Gini coefficient for income rose from 0.48 in 1982 to 0.57 in 2006 (Wolff, 2014, p. 27). Furthermore, the share of national income held by the richest 1% in the USA increased from8% to 18.9% between 1980 and 2012 (Alvaredo, Atkinson, Piketty & Saez, 2011). Growth in inequality at the top tail of the distribution was driven by the financial sector, with financial services employees accounting for 15%-27% of the top 0.1% of the income distribution in the USA, compared to 6% by the non-financial sector executives (Kaplan & Rauh, 2009). Simultaneously, due to wage growth lagging behind productivity growth, the share of worker compensation in GDP declined steadily from 62% in 1980 to 56% in 2013 in the USA (AMECO Database), suggesting redistribution

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of national income towards profits (and more specifically financial profits).

Numerous studies have attempted to formally link financial sector transformation with rising income inequality. Van Treeck and Sturn (2012), Mian and Sufi (2013), Cynamon and Fazzari (2014), Stockhammer (2015) argue that increasing inequality levels were the key contributing force to the Great Recession. In this paper, the focus is on the impact of financial sector transformation on income distribution. A few econometric studies establish a positive association between the two (Assa, 2012; Kus, 2012; Rosnick & Baker, 2012; Arestis, Charles & Fontana, 2013; Jerzmanowski & Nabar, 2013; Lin & Tomaskovic-Devey, 2013; Van Arnum & Naples, 2013). Inequality is generally proxied by the Gini coefficient or the labour share of income, while financial sector transformation tends to be measured as the relative size of the financial sector, stock market or bank income to GDP.

The purpose of this paper is to revisit the relationship between finance and inequality in the light of newly available data on income concentration in advanced countries. We aim to test the association between different measures of financial sector transformation, reflecting potential transmission mechanisms generating inequality, and the share of national income going to the top 10% of the population. The structure of this paper is as follows. Section II presents method, data sources and definitions underlying our empirical specification. In Section III, results of the econometric estimation are analysed. Section IV discusses the robustness of results and issues remaining for future research. Overall, specific aspects of financial sector transformation are found to contribute to income concentration, although the interpretation is riddled with problems regarding conceptual framework, measurement and empirical specification.

Data and method

We develop an econometric model describing the influence of financial sector transformation on income distribution, employing annual data for 16 OECD countries[1] between 1995 and 2009.

The dependent variable in the model measures the share of national income flowing to the top 10% of the population. Since there is no data on the shape of the whole cumulative income distribution for all countries in our panel, the focus is placed on the concentration of income in the top decile. We deem it more appropriate for the purposes of the model than other measures of inequality found in the literature like the Gini coefficient and the wage share. The wage share excludes those outside the labour market who may be impacted by the financialisation processes. Moreover, it disguises the heterogeneity of earnings as large salaries of top managers and wages of non-managerial workers are captured in one measure. In turn, the Gini is sensitive to how income is classified in different datasets (Atkinson & Brandolini, 2001, p. 781). Furthermore, since the Gini is a relative measure, two countries with distributions differing in absolute terms may have the same Gini coefficients. It is also sensitive to the unit of analysis (individuals versus households) and underlying definitions (Deininger & Squire, 1996). Moreover, transfers between different levels of distribution may not have equal weight in changing the coefficient value (Cowell, 2011, p. 26).

Moreover, the Gini is too aggregate a measure to provide meaningful insight into the transmission mechanisms through which financial sector transformation has influenced income distribution. This is because many of the channels of influence may not be observable directly at such a level of aggregation as inequality itself is an outcome of the distributional shifts. By focusing on the income share we are better able to understand which aspects of financialisation are associated with the income shift towards to top earners. In contrast, the top decile income share avoids the problematic assumption regarding weighting of transfers across the distribution and the sensitivity of Gini to income and unit definitions

Data on the top 10% income share are obtained from the World Top Incomes Database by Alvaredo et al. (2011) for a representative group of 16 OECD countries, characterised by different institutional arrangements and hence a varying depth of financial sector transformation processes. This data is collected based on income tax reports in each country and presents gross nominal income held by the top decile of the population relative to national income. The most serious problem we encounter in using this measure is underreporting of the richest, who are thereby trying to minimise the amount paid in taxes. Alvaredo et al. deal with this problem by applying Pareto interpolation to approximate the top tail

1 Australia, Canada, Denmark, Finland, France, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, United Kingdom, USA.

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of the distribution (Atkinson et al., 2011, p. 14). However, in the recent years much of the income of the richest has become "hidden" in tax havens outside of a given country's accounting system and has been held in various mutual and hedge funds and other investment vehicles. Not only are these assets taxed at much lower rates or not taxed at all but also data on these types of funds are not consistently reported by official statistics in any country. Consequently, the top 10% share reported by Alvaredo et al. likely underestimate the true share of the richest 10% in national income.

Since our goal is to provide a comprehensive picture of the distributive effects of financial sector transformation, we do not attempt to capture them in one indicator. We separately estimate three different measures of financial sector transformation encountered in the empirical work, corresponding to various transmission channels through which financialisation may have impacted inequality. These include the GDP share of the total value traded in stock markets, the value of private credit relative to GDP and the share of pre-tax bank income in GDP. The former two indicators are collected from the financial structure database by Beck, Demirgüç-Kunt and Levine (2000) while the latter comes from the OECD Financial Accounts. The relative sizes of the financial sector value added and employment are intentionally neglected as they carry little insight into transmission mechanisms of redistribution due to their high level of aggregation.

Firstly, the total value traded in stock markets is defined as the value of all shares traded in the stock market relative to GDP. We choose this measure overstock market capitalisation or turnover as it depicts the expansion of trading in stock markets and thus proxies the increased involvement of economic agents in the financial sector, reflecting greater reliance on the external sources of financing. Due to insufficient data we are not able to account for the total value traded in the bond market. We expect that greater volume traded in stock markets increases the top decile income share as participation in stock markets and the resulting income flows are more likely among the top population percentiles.

Secondly, transformation of the banking sector is considered one of the most pronounced channels of distribution (cf. Kus, 2012; Van Arnum & Naples, 2013). We resort to the GDP share of pre-tax income reported by domestic banks to proxy the banking sector expansion, remembering the potential downward bias associated with income underreporting. The percentage share of bank pre-tax income to GDP measures profitability of the banking sector arising from market activity before reduction by tax payments. Expansion of banks' pre-tax profits has been one of the most manifest processes of financial sector transformation benefitting only a narrow group of managers and financial investors. Hence, it is expected to contribute to greater share of income flowing to the top 10%. The effect is likely to be particularly important in the context of weakening the traditional intermediary role of banks as fee income obtained from security underwriting tends to exceed the conventional interest revenue.

Thirdly, the private credit share of GDP is defined as claims of both commercial banks and non-bank financial institutions on the private sector as a percentage of GDP. Although far from adequate, it is the only available indicator allowing us to capture the role of non-bank intermediaries. The distributive effect of private credit expansion is ambiguous. On the one hand, it may be equalising if it is extended to those needing credit to sustain consumption but it may have the opposite effect if it occurs in result of greater securitisation. In the latter case, it is the richest that are more likely to benefit from a greater amount of credit for financial investment. The data for bank income and private credit share are consistently available only until 2008. Insufficient observations force us to exclude Australia, Ireland and the UK from the bank income regression as well as Norway from the private credit model.

To isolate the effect of financial sector transformation on the top decile's income share, we control for other forces contributing to greater income concentration which have been consistently accounted for in the reviewed literature. One of the most important factors is the decline in bargaining power of the working class resulting from the erosion of trade unions. This shift in labour market institutions is measured by union density, i.e. the ratio of workers belonging to a trade union to all workers in the economy, collected from ICTWSS data. Despite some criticism of the use of union density to measure labour militancy (cf. Visser, 2006; Howell, Baker, Glyn & Schmitt, 2007), it provides a good depiction of the penetration of trade unions into the workforce. Greater union density is expected to lower the share of income going to the top decile.

Secondly, to investigate the degree to which financial

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sector transformation influences growth of the top incomes beyond general economic growth, we control for the annual growth rate of GDP per capita from OECD National Accounts. Thirdly, we control for macroeconomic conditions, proxied by the harmonised unemployment rate (OECD). Greater unemployment, signifying lower labour demand, directly contributes to the redistribution of income as the unemployed lose a stable source of income. Consequently, unemployment is expected to be positively associated with income concentration at the top decile.

In addition to these standard controls, we propose to control for property price inflation. This is because financial sector growth in the 2000s relied on a housing price bubble as mortgages constituted the basis of most securitised products (cf. Pollin & Heintz, 2013). There are strong reasons to argue that property price growth had an impact on inequality levels, albeit the exact empirical effect is not clear. Since mortgages were held primarily by households at the bottom and middle of the income distribution, growth in house prices was more equalising in so far as it boosted the asset side of household balance sheets and allowed for more equity withdrawal from housing acting as a collateral for further loans. However, as securities based on those mortgages were owned by wealthy financial investors, reversal of the housing bubble since 2006 increased wealth inequality (Wolff, 2014, p. 27). This is because mortgage-based securities were classified as more senior than the underlying mortgages so that the flow of cash from mortgages to the senior securitised tranches was guaranteed irrespective of the repayment capacities of the mortgage-holders. Consequently, property price deflation was fatal for the solvency of the latter group as the real value of debt increased (Fisher, 1933), contributing to growing wealth inequality since the crisis. Data for the annual growth rate of real housing prices are collected from the OECD Analytical House Price Database.

Furthermore, two controls are employed to account for institutional differences between countries in our sample. First is the openness of the capital account, measured as the value of total external assets and liabilities to GDP, using a database by Lane and Milesi-Ferretti (2007). It is expected that the indicator is positively associated with the top 10% income share as global financial investment tends to be concentrated in a narrow group of the most affluent investors. Secondly, we account for the magnitude of social transfers, corresponding to institutional differences between countries in our sample. We proxy the depth of social policy by the percentage share of social spending in GDP by the central government obtained from OECD National Accounts. *Ceteris paribus*, we expect it to have an equalising impact on the distribution and decrease the income share of the top 10%.

Equation 1 presents the baseline specification, where i and t correspond to country and time variables in the panel. Y represents the top 10% income share, X is each of the chosen measures of financial sector transformation(each regressed individually) and Wj corresponds to the control variables.

$$Y_{i,t} = \beta_0 + \beta_1 X_{i,t} + \beta_j W_{j,i,t} + \varepsilon \qquad (1)$$

The mean top 10% income share in our sample is relatively high at 33%. Mean total value traded in stock market constitutes 89.7% of GDP, while mean bank income and private credit shares are 1.5% and 118.7% of GDP respectively. Among the control variables, mean union density is 34.8% of workforce. In the macroeconomic conditions, annual GDP *per capita* growth rate is 1.6%, unemployment rate is on average 6.8% and mean real housing price growth rate is 3.8%. Mean social spending share of GDP is 21.5%. The mean value of total external assets and liabilities to GDP measuring capital account openness is 432.2%. A correlation test reveals that variables in our models tend to be slightly correlated, which is understandable due to the interconnectedness of macroeconomic indicators. Consequently, a multicollinearity problem should be present but not severe in our estimation.

Based on the results of specification tests (see appendix), fixed-effects regressions are estimated for all models. By using fixed effects regressions, it is assumed that individual characteristics of a country in our panel affect estimates (Greene, 2003, p. 287). Further diagnostic tests indicate the presence of non-spherical errors in the form of both heteroscedasticity and autocorrelation, suggesting that estimated standard errors for the coefficients are inflated and thus unreliable in predicting statistical significance of the estimates. To account for this problem, we use Driscoll-Kraay panel-corrected standard errors. While such correction does not completely remove the bias of non-spherical standard errors, it allows for determination of more reliable confidence intervals and thus statistical significance without assumptions about

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the form of heteroscedasticity (Driscoll & Kraay, 1998). The next sections present our empirical results and discuss the emerging estimation problems.

Results

Table 1 presents estimation results. For the stock market regression, *ceteris paribus*, one percentage point increase in the GDP share of the total value traded in stock market is associated with 0.01 percentage point increase in the top 10% share. The estimates of stock market expansion are highly significant at 1%, so that participation in stock markets is found to benefit the top population decile. The coefficient of the pre-tax bank income share of GDP is found to be negative and significant at 1% level. All else equal one-unit increase in the share of bank income in GDP is associated with 0.14 percentage point fall in the share of income flowing to the top decile. The estimate of the GDP share of private credit issued by bank- and non-bank intermediaries is negative but insignificant in the baseline fixed-effects specification. *Ceteris paribus*, one-unit rise in the relative size of private credit to GDP is associated with 0.002 percentage point fall in the top 10% income share.

Among control variables, union density emerges as one of the most powerful channels of redistribution. Its association with the top 10% income share is consistently negative and highly significant at 1% level. The coefficient ranges from -0.2 to -0.3, implying that higher union density counteracts the concentration of income at the top. The importance of unionisation in our model arises as a suppressed labour voice transfers the bargaining power to the top earners. While we do not explicitly model the size of the wage premium, we suspect that weaker labour militancy allows those occupying top managerial and supervisory positions to capture a greater share of earnings relative to the rest of the working class.

Another robust variable is the annual growth of per capita GDP, whose coefficient is positive and significant at 1% level in all specifications. *Ceteris paribus*, one-percentage point faster growth of per-capita GDP is associated with 0.2-0.3 unit increase in the top 10% income share. In terms of macroeconomic conditions, unemployment rate estimate is positive but of varying significance, ranging from 0.03 to 0.2. It is significant at 5% level in the baseline model of bank income, where *ceteris paribus* one-unit rise in unemployment rate is associated with around 0.08 percentage point increase in income concentration at the top decile, as well as in the stock market regression, where the estimate of 0.2 is significant at 1% level.

Surprisingly, the share of social spending in GDP is persistently found to contribute to greater income concentration, with a positive and statistically significant coefficient in all models. *Ceteris paribus*, one-unit rise in the GDP share of government's social spending is associated with 0.3-0.4 unit increase in the share of income going to the top decile. The sign of the social spending coefficient is contrary to our initial expectations. This may be because our measure does not gauge the composition of social expenditure. It seems plausible that certain types of transfers may not be sufficient to compensate for the concentration of market income and suggests that fiscal policy targets may be inadequately formulated.

The estimate of capital account openness is positive although of varying significance. It is significant in the regression of bank income but not in the model for stock market value traded or private credit. When significant, a one-unit increase in the relative size of total external assets and liabilities to GDP is associated with 0.004 unit rise in the top 10% income share. This may be because international markets are more accessible to the top earners and are an important source of financial investment through which income and thus savings of the richest are held.

The coefficient of real housing price growth rate is negative but insignificant in most of the specifications. In the regression of private credit it is significant at 1% level, implying that one unit rise in the rate of house price growth is associated with a 0.02 percentage point decrease in the top 10% income share. A negative estimate suggests that housing price inflation had an equalising effect. This is due to greater possibilities for housing equity withdrawal, boosting the balance sheets for households along the distribution. In the baseline regression of stock market value traded, the coefficient of house prices growth rate becomes positive but is not significant.

Goodness-of-fit of the models measured by within R^2, i.e. the proportion of variation in the dependent variable explained by the regressors ignoring fixed effects, is moderate for a panel regression, ranging 0.3-0.4 depending on specification. Reported R^2 are not comparable across different estimations as no adjustment is made for the default inflation of the coefficient on

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Table 1: Results of the fixed effects regressions

Model / Variable	Stock Market Total Value Traded Regression	Bank Income Regression	Private Credit Regression
Stock Market Total Value Traded to GDP	.011*** (6.78)	-	-
Bank Income to GDP	-	-.138*** (-3.75)	-
Private Credit to GDP	-	-	-0,002 (-0.74)
Union Density	-.226*** (-14.71)	-.325*** (-10.70)	-.315*** (-14.27)
GDP per capita growth rate	.167** (2.51)	.214*** (2.62)	.299*** (7.77)
Unemployment Rate	.166*** (11.39)	.078** (2.27)	0,026 (1.36)
Government Social Expenditure to GDP	.254*** (5.75)	.340*** (3.44)	.394*** (7.44)
Capital account openness	0 (0.69)	.004*** (11.03)	0,001 (1.16)
Real house price growth rate	0,017 (0.83)	-0,007 (-0.44)	-.022*** (-2.62)
Constant	32.899*** (25.23)	34.959*** (17.39)	34.885*** (30.67)
Within R^2	0,405	0,392	0,399
Number of observations	240	182	210
Period of analysis	1995-2009	1995-2008	1995-2008
Countries excluded	-	Australia, Ireland, UK	Norway

Source: OECD Database, Beck et al. (2000), ICTWSS, Lane&Milesi-Ferretti (2007)

*Notes: *** = 1% statisticalsignificance of two-tailed test; ** = 5% significance; *=10% significance*

inclusion of additional variables.

Discussion

One of the problems of the estimated model is the issue of endogeneity arising due to simultaneous causality between our dependent and independent variables. It is likely that the chosen measures of financial sector transformation, i.e. the shares of stock market activity, bank income and private credit in GDP can be larger owing to greater investment demand and deposits among the richest 10%. For this reason, we conduct a sensitivity analysis using the Arellano-Bover/Blundell-Bond difference Generalised Method of Moments (GMM) estimation in order to check the robustness of the obtained estimates.

GMM is a dynamic panel data model that deals with endogeneity by utilising information on the past values of the endogenous variables. Difference GMM addresses endogeneity by transforming endogenous variables by first differencing, assuming that the emergent instruments are uncorrelated with the panel fixed effects (Roodman, 2009, p. 86). Consequently, no intercept is reported. This method is particularly suitable in panels with short time series but extensive cross-sectional dimensions, like our sample (ibid.).

Re-estimation of the model equations using GMM reveals that some of the estimates are volatile to bias

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induced by the presence of endogeneity (Table 2). The estimate of stock market value traded loses its high significance, retaining it only at 10% level, while the estimate of bank income becomes not significant. Moreover, the magnitude of the coefficient decreases from 0.01 to 0.004. However, in the case of private credit, GMM estimation yields a larger effect significant at 10% level. Furthermore, in the stock market regression government expenditure becomes insignificant and negative, while capital account openness becomes significant at 1% level with a positive coefficient of 0.001. However, the variable loses significance in the bank income regression. Similarly, real property price growth becomes insignificant in the private credit model. Consequently, there is evidence of a certain endogeneity bias in the fixed-effects specification. Notably, union density and GDP per capita growth rate emerge as the most robust in terms of significance and estimate size.

Overall, our results show that stock market expansion is the most robust transmission mechanism of financial sector transformation towards national income concentration at the top population decile. The sensitivity of the expansion in bank profitability and private credit estimates to specification and estimation method can be explained by measurement problems and omitted variable bias. At such a level of aggregation, varying accounting practices and institutional characteristics of countries in our panel are likely to be captured by the regression's error term, resulting in different patterns of interaction with the estimates in each specification. The fixed effects model seems incapable of controlling for all the heterogeneity between entities. Moreover, aspects of financialisation omitted due to the lack of explicit data may be implicitly present in the error term and simultaneously

Table 2: Sensitivity analysis – General Method of Moments

Model Variable	Stock Market Total Value Traded Regression	Bank Income Regression	Private Credit Regression
Stock Market Total Value Traded to GDP	.004* (1.88)	-	-
Bank Income to GDP	-	-.046 (-0.65)	-
Private Credit to GDP	-	-	-.011** (-2.35)
Union Density	-.356*** (-6.53)	-.441*** (-7.04)	-.487*** (-10.46)
GDP per capita growth rate	.150*** (3.71)	.206*** (4.50)	.247*** (6.06)
Unemployment Rate	.185*** (3.26)	0,09 (1.50)	0,042 (0.69)
Government Social Expenditure to GDP	-0,005 (-0.06)	.175** (2.17)	.300*** (4.23)
Capital account openness	.001*** (2.72)	0,001 (0.53)	0 (0.90)
Real house price growth rate	-0,011 (-0.76)	0 (-0.01)	-0,019 (-1.51)
Number of observations	208	168	180
Period of analysis	1995-2009	1995-2008	1995-2008
Countries excluded	-	Australia, Ireland, UK	Norway

Source: OECD Database, Beck et al. (2000), ICTWSS, Lane & Milesi-Ferretti (2007)

*Notes: *** = 1% statistical significance of two-tailed test; ** = 5% significance; *=10% significance*

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correlated with the included financialisation variables. Since the financial sector transformation processes presented in our model in both key financialisation regressors and financialisation controls interact with each other and other economic variables, bias of the estimates is unavoidable, leading to volatile and unexpected results.

Further limitation of our approach is a linear treatment of the relationships. Since our aim is to maintain a dialogue with the literature, the possibility of non-linear estimation remains unexplored. Given the complexity of economic phenomenon, it is likely that some of the distributive forces of financialisation do not have constant dynamics. The sensitivity of our model to temporal dynamics is signalled by the augmented Dickey-Fuller test for unit root. It indicates that the top 10% income share is non-stationary, signifying that its mean is not constant overtime but follows a stochastic trend, which can undermine the reliability of our results.

Finally, the model does not directly test the causality of financial sector transformation on increasing the top 10% share. It remains riddled with problems of endogeneity and GMM correction may not eliminate all of the bias, particularly in the case of bank profitability and private credit expansion. These problems expose the limitations of empirical methods faced with complex and interconnected economic phenomena such as financial sector transformation.

Conclusion

This paper argues that the expansion of financial markets in economic policy and decision-making observed since the 1980s is a highly complex phenomenon creating forces of distribution of income towards the richest. We find that out of three chosen measures of financial sector transformation, the size of the stock market relative to GDP is the most significant aspect associated within, come concentration at the top population decile. Union density and growth rate of per capita GDP are the most robust variables associated with income concentration at the top of the distribution. Surprisingly, social expenditure is found to contribute positively to income inequality in the sample over the period estimated. The results, however, are subject to various estimation problems, which remain only partially resolved.

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Technical appendix

1. Choice of specification, fixed versus random effects: Hausman test.

Note: Where Hausman testwas not conclusive, the fixed effects model was preferred to maintain consistency with the remainder of the estimations

a) Stock market total value traded %GDP

---- Coefficients ----

	(b) fe_stockm	(B) re_stockm	(b-B) Difference	sqrt(diag(V_b-V_B)) S.E.
StockMarketGDP	.012122	.0126209	-.0004989	.000296
UnionDensity	-.2260643	-.1881363	-.037928	.0276942
UnemploymentRate	.166392	.1594567	.0069352	.0100845
GDPpcgrowth	.166508	.1483662	.0181417	.0058423
SocialExpendGDP	.253732	.2155137	.0382183	.0122692
CapitalAccOpenness	.0003714	.0005592	-.0001878	.0001355
RealHousePrGrowth	.0165733	.0166361	-.0000629	.

b = consistent under Ho and Ha; obtained from xtreg

B = inconsistent under Ha, efficient under Ho; obtained from xtreg

Test: Ho: difference in coefficients not systematic

chi2(7) = (b-B)'[(V_b-V_B)^(-1)](b-B)
= 12.07
Prob>chi2 = 0.0982
(V_b-V_B is not positive definite)

b) Bank income %GDP

---- Coefficients ----

	(b) re_b	(B) fe_b	(b-B) Difference	sqrt(diag(V_b-V_B)) S.E.
BankIncomeGDP	-.123462	-.1381039	.0146419	.0380035
UnionDensity	-.2105484	-.3250059	.1144576	.
UnemploymentRate	.0474829	.0782162	-.0307333	.0192375
GDPpcgrowth	.1748802	.213523	-.0386428	.0251672
SocialExpendGDP	.2385611	.3399406	-.1013795	.0163662
CapitalAccOpenness	.0041294	.0036493	.0004802	.
RealHousePrGrowth	-.0093248	-.0067473	-.0025775	.0076347

b = consistent under Ho and Ha; obtained from xtreg

B = inconsistent under Ha, efficient under Ho; obtained from xtreg

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Test: Ho: difference in coefficients not systematic

chi2(7) = (b-B)'[(V_b-V_B)^(-1)](b-B)
= 77.58
Prob>chi2 = 0.0000
(V_b-V_B is not positive definite)

Choose Fixed Effects

c) Private credit %GDP

---- Coefficients ----

	(b) fe_credit	(B) re_credit	(b-B) Difference	sqrt(diag(V_b-V_B)) S.E.
PrivateCreditGDP	-.0020045	.0001154	-.0021199	.
UnionDensity	-.3150916	-.2452726	-.0698189	.0261574
UnemploymentRate	.0263327	.0119399	.0143929	.
GDPpcgrowth	.2994005	.2876206	.0117799	.
SocialExpendGDP	.3939122	.3390166	.0548956	.0067319
CapitalAccOpenness	.0007381	.0010788	-.0003406	.0001199
RealHousePrGrowth	-.0222606	-.0208126	-.001448	.

b = consistent under Ho and Ha; obtained from xtreg

B = inconsistent under Ha, efficient under Ho; obtained from xtreg

Test: Ho: difference in coefficients not systematic

chi2(7) = (b-B)'[(V_b-V_B)^(-1)](b-B)
= 66.01
Prob>chi2 = 0.0000
(V_b-V_B is not positive definite)

Choose Fixed Effects

2. Modified Wald test for groupwise heteroskedasticity in fixed effect regression model

Regression	**H0: no heteroscedasticity**	
Stock market total value traded %GDP	*chi2(16)*	*Prob>chi2*
	2170.57	0.0000
Bank Income %GDP	*chi2(13)*	*Prob>chi2*
	1224.93	0.0000
Private credit %GDP	*chi2(15)*	*Prob>chi2*
	573.31	0.0000

In all cases H_0 of no heteroscedasticity is rejected.

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3. Woodridge test for autocorrelation in panel data

Regression	**H_0: no first-order autocorrelation**	
Stock market total value traded %GDP	*F(1, 15)*	*Prob>F*
	7.234	0.0168
Bank Income %GDP	*F(1, 12)*	*Prob>F*
	6.719	0.0236
Private credit %GDP	*F(1, 14)*	*Prob>F*
	12.721	0.0031

In all cases H_0 of no autocorrelation is rejected.

4. Non-stationarity

Fisher-type unit-root test for Top10share

Based on augmented Dickey-Fuller tests

Ho: All panels contain unit roots	Number of panels	=	16
Ha: At least one panel is stationary	Avg. number of periods	=	19.81

Statistic		p-value	
Inverse chi-squared(32)	P	33.5631	0.3915
Inverse normal	Z	0.0017	0.5007
Inverse logit t(84)	L*	-0.0390	0.4845
Modified inv. chi-squared	Pm	0.1954	0.4225

H0 not rejected – unit root present

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