Online Appendix

This online appendix presents additional results which did not find their way into the paper due to space constraints.

Granger Causality Tests

The table below presents the results for bivariate Granger causality tests between our household debt measure (DH) and real property price indices (PP). We find that property prices Granger-cause household debt independent of the lag structure chosen. In contrast the findings for Granger causality from household debt to property prices is sensitive to the lag structure: we fail to find evidence for Granger causality with 1 lag but find evidence with two lags. Thus, we find more robust evidence in support of Granger causality running from property prices towards debt than in the other direction. Theoretically one can justify a unidirectional relationship running from property prices to household debt based on the notions of speculative dynamics in real estate markets as documented by Shiller (2015) and formally expressed by momentum vs fundamentals trading models (Dieci and Westerhoff 2016). We discuss this literature in section 2 of the paper. If such speculative dynamics in the housing market are present, the price expectations of momentum and fundamentalist traders rather than household liabilities are the driving factor of real estate price dynamics. Liabilities are passively dragged along in this case. We do think that there is indeed a strong speculative element present in housing markets which provides a justification of our empirical approach.

null hypothesis	PP does not g-causes D		D does not g-causes PP	
dep var	DH		PP	
specification	1 lag	2 lags	1 lag	2 lags
log(DH-1)	0.733***	0.931***	0.103	0.333***
	-0.04	-0.05	-0.07	-0.1
log(DH-2)		-0.204***		-0.129
		-0.05		-0.09
log(PP-1)	0.098***	0.112***	0.577***	0.698***
	-0.03	-0.03	-0.05	-0.05
log(PP-2)		-0.054*		-0.281***
		-0.03		-0.05
_cons	0.010***	0.011***	0.002	0.001
	0	0	0	0
N	460	442	460	442
H0: PP=0	0.00	0.00		
H0: D=0			0.12	0.00

Single Country Results

In addition to the panel error correction models for which results are reported in the paper, we also estimated single country EC models for the UK and the US. We did this in order to test whether our findings are driven by the homogeneity assumption implicit in the panel approach and whether those countries which experienced especially strong shifts in the distribution of income like the US or the UK provide evidence consistent with the expenditure cascades hypothesis. However the table reported below reveals that in neither of these two countries we find a statistically significant and positive long-run impact of an increasingly polarized income distribution and outstanding household sector liabilities. Thus, our results are not driven or invalidated by the pooling assumption we have made.

	US	UK
long run		
log(YD)	1.733***	0.755***
	(0.16)	(0.07)
Top1	-4.984***	-0.357
	(1.79)	(0.81)
log(PP)	0.696***	0.827***
	(0.12)	(0.07)
R	3.318**	4.851***
	(1.41)	(1.18)
OLD	15.367***	4.497**
	(3.44)	(2.25)
log(CRED)	-0.143	0.209**
	(0.13)	(0.10)
short run		
adjustment	-0.316***	-0.373***
	(0.07)	(0.03)
N	30	24
AR(1) p-val.	0.10	0.03
AR(2) p-val.	0.13	0.04
AR(3) p-val.	0.09	0.04

Adjusted Wage Share as Distributional Variable

The table below reports the results when the adjusted wage share (WS) is used as the income distribution measure in contrast to the top income share or the Gini coefficient. This serves as an additional robustness check of our conclusion that polarizations in the distribution of income are not the main drivers of household sector indebtedness. As can be seen from the table below neither the DFE nor the PMG estimator yield statistically significant and positive long run wage share coefficients.

	DFE	PMG
long run		
log(YD)	0.888***	0.699***
	(0.20)	(0.04)
WS	-1.117	-1.250***
	(0.93)	(0.14)
log(PP)	0.475***	0.708***
	(0.14)	(0.02)
R	-3.089**	0.667***
	(1.46)	(0.21)
OLD	0.625	1.350*
	(1.64)	(0.69)
log(CRED)	0.580**	0.181***
	(0.27)	(0.05)
short run		
adjustment	-0.059***	-0.122***
	(0.01)	(0.04)
$\Delta\log(Y^D)$	0.155***	
	(0.06)	
Δlog(PP)	0.208***	0.206***
	(0.02)	(0.05)
ΔR	0.174***	
	(0.06)	
ΔOLD	-1.367**	
	(0.63)	
Δlog(CRED)	-0.068***	-0.103***
	(0.03)	(0.03)
$\Delta\log(DH_{t-1})$	0.679***	0.523***
	(0.03)	(0.07)
$\Delta\log(Y_{t-1}^D)$	-0.117**	
	(0.06)	
$\Delta\log(PP_{t-1})$	-0.101***	-0.114**
	(0.03)	(0.06)
constant	0.024	0.307***
	(0.11)	(0.10)
N	380	380

Confidence Bounds Around Contribution of Top Income Share

We have carried out the additional tests to see whether cumulative inequality effects based on the upper bound of the estimated coefficients' confidence bands produce results which would justify a re-assessment of our conclusions. The results are presented in the table below:

	(1) actual Δ in D/Y^D	(2) lower bound pred. Δ in D/Y^D	(3) upper bound pred. Δ in D/Y^D	(4) $TOP1$ lower bound	(5) $TOP1$ upper bound
(A) DFE - TOP1	54%	38%	65%	-10%	8%
(B) PMG - TOP1	54%	42%	52%	-2%	5%

The table displays the total change in debt-to-income ratios the model predicts (based on the estimated coefficients in Table 4) if one uses the upper and lower bounds of the Top 1% income share coefficient (columns 2 and 3) in comparison with the actual change in debt to income ratios (column 1). Upper and lower bounds are the point estimate plus/minus two standard errors. The point estimates of the Top 1% coefficient are -0.67 and 0.45 for the DFE and PMG estimator respectively (see Table 4 in the paper). The two standard error upper bounds are 3.1 and 1.96 respectively.

Columns 4 and 5 report the individual contribution of the top income share based on the upper/lower bound of the coefficient estimate. One can see that even if one uses the upper bounds, the contribution of the top income share is clearly second to the property price contribution which amounts to 25% and 36% in the DFE and PMG specifications respectively.

So, while non-trivial contributions of the top income share lie within a two standard error confidence band, our conclusion that property prices are the more important factor is not changed.