Unequal Multipliers: Income Inequality and Government Spending*

Donggyu Lee†
Federal Reserve Bank of New York

Choongryul Yang‡
Federal Reserve Bank of Atlanta

Abstract

Using U.S. state-level military spending and top income shares for 1966–2006, we study how local fiscal multipliers vary with income inequality. Output multipliers are smaller in high-inequality states, while employment multipliers do not differ significantly. A decomposition accounts for the smaller output response through output per worker and income estimates show smaller personal-income gains in high-inequality states. We interpret these results using a two-agent New Keynesian model with capital–skill complementarity and segmented labor markets. In the calibrated model, a shift in factor demand toward capital and high-skill labor raises income concentration and lowers the income reaching hand-to-mouth households per spending dollar. This weakens the feedback from income to private demand. Adjustment through wages, hiring, and hours, together with a different composition of employment, gives the high-inequality economy a slightly larger employment response despite its smaller output multiplier. The model links inequality to fiscal transmission through production structure and income incidence even when the hand-to-mouth population share is unchanged.

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†33 Liberty Street, New York, NY 10045, U.S.A. Email: donggyu.lee@ny.frb.org

‡1000 Peachtree Street NE, Atlanta, GA 30309, U.S.A. Email: choongryul.yang@atl.frb.org

1 Introduction

How does the income distribution shape the effectiveness of government spending? Households differ in their marginal propensities to consume (MPCs), so fiscal transmission depends on both the income generated and its recipients. For a given income gain, consumption rises more when a larger share reaches households with high MPCs (Galí, López-Salido and Vallés, 2007, Kaplan, Moll and Violante, 2018). Higher top income concentration need not imply a larger share of financially constrained households or a larger fiscal multiplier. We study how local output and employment multipliers vary with income inequality, and examine the production and income responses behind that relationship.

We use U.S. state-level military spending and top income shares for 1966–2006. The instrumental variable design of Nakamura and Steinsson (2014) exploits differences across states in the response of military spending to national buildups and drawdowns. We extend this design by interacting spending with lagged top income shares from the Frank-Sommeiller-Price series (Frank, Sommeiller, Price and Saez, 2015). The comparison concerns relative state responses under common national policy. Its interpretation requires that the instruments be unrelated to other determinants of relative state outcomes. Year fixed effects absorb common national changes, although they need not absorb differences in states' responses to those changes.

Output multipliers are smaller in states with greater top income concentration. A one percentage point increase in the lagged top 10% income share relative to the cross-state median is associated with an output multiplier that is 0.19 lower. The employment interaction is positive but statistically insignificant in the baseline specification. Local projections show that the output interaction remains negative and significant through horizon three. The employment interaction becomes positive and significant at later horizons. The initial output-employment contrast therefore concerns the baseline response, rather than equality of employment responses at every horizon.

We first examine this contrast through an accounting decomposition. On a common sample, the negative output interaction is accounted for by a smaller response of output per worker, partly offset by the positive employment point estimate. This finding directs attention to the production associated with the employment response.

The income accounts then show which income gains accompany the smaller output response in high-inequality states. Using the Bureau of Economic Analysis (BEA) state personal income accounts, we estimate the income generated per dollar of military spending. Personal income has point estimates of \$1.62 in a state five percentage points below the median top 10% share and \$0.54

in a state five points above. Earnings by place of work account for most of this difference in point estimates, with responses of $1.67 and $0.68.

Observed state characteristics provide further comparisons. Adding a government spending interaction with the earnings share attenuates the inequality coefficient by 39%, and adding union density attenuates it by 27%. The earnings share divides earnings by earnings plus dividends, interest, and rent. Compensation divided by GDP does not produce the same attenuation, and the CPS bottom-to-median income ratio does not account for the inequality interaction. The income accounts also do not directly measure high-MPC households' receipts, and skill-specific income responses vary across periods. The evidence therefore leaves open whether the smaller output multiplier in high-inequality states can arise from differences in the income reaching high-MPC households, even when their population share is unchanged.

To examine this possibility, we use a two-agent New Keynesian model with capital–skill complementarity and segmented labor markets. A controlled experiment reallocates incremental wage income from hand-to-mouth (HtM) households to Ricardian households while retaining the same steady state and production structure. The output multiplier falls because less of the additional income is spent on current consumption. We then consider a shift in factor demand toward capital and high-skill labor. In the calibrated model, this raises top income concentration and lowers income reaching HtM households per spending dollar, even with their population share held fixed. Wages, hiring, and hours adjust, and the composition of employment differs, so that output rises less while the headcount employment response is slightly higher in the high-inequality economy.

The model separates the total household-income gain per spending dollar from the share of that gain received by HtM households. Their product is the income reaching HtM households per spending dollar. We calibrate the model to the output and employment multipliers and their inequality interactions, so the fit to these four moments is part of the calibration. We then use the model to show how production structure and labor-market frictions generate the contrast between output and employment multipliers that we find in the data.

Our empirical contribution concerns a dimension of fiscal transmission that differs from the macroeconomic conditions emphasized in much of the multiplier literature. Previous work studies the business cycle (Auerbach and Gorodnichenko, 2013, 2012, Ramey and Zubairy, 2018), the exchange rate regime and public debt (Ilzetzki, Mendoza and Végh, 2013), the zero lower bound (Christiano, Eichenbaum and Rebelo, 2011), and tax financing (Ferriere and Navarro, 2025). Cloyne, Jordà and Taylor (2023) develop methods for estimating state-dependent local projections. We document how relative state responses vary with the income distribution. The output interaction

remains negative when we control for state tax progressivity, although that estimate is imprecise in the shorter sample.

The distinction between income gains and recipients connects the evidence to the literature on two-agent New Keynesian (TANK) and heterogeneous-agent New Keynesian (HANK) models. Auclert (2019) shows how the distributional effects of monetary policy contribute to aggregate consumption responses when households differ in their MPCs. Bilbiie (2020) emphasizes how HtM income responds to aggregate activity in the New Keynesian cross. The intertemporal Keynesian cross extends the analysis to heterogeneous households and dynamic responses (Auclert, Rognlie and Straub, 2024). Broer, Krusell and Öberg (2023) show how the distribution of wage and profit income interacts with wage and price rigidities to determine fiscal multipliers. These frameworks motivate examining the income generated by spending and its recipients (Auclert, Rognlie and Straub, 2025). Our contribution combines estimates by income source with a model in which production structure determines how additional income reaches households with different MPCs.

The findings also relate to the link between inequality and aggregate demand. Summers (2015) identifies rising inequality as a source of higher desired saving that can weaken aggregate demand. Auerbach, Gorodnichenko and Murphy (2021) formalize a demand-determined economy in which inequality reduces output and dampens fiscal stimulus. Auclert and Rognlie (2020) find small short-run output effects of greater income inequality, with larger long-run effects when inequality reflects greater income risk. We examine income incidence as a possible source of the differences in local output responses between low- and high-inequality states.

Related work finds different relationships between inequality and fiscal and monetary transmission.[1] Brinca, Holter, Krusell and Malafry (2016) find larger fiscal multipliers in countries with greater wealth inequality, using a cross-country structural VAR that assumes government consumption cannot respond to output within the quarter. Brinca, Ferreira, Franco, Holter and Malafry (2021) document larger output losses from fiscal consolidation in countries with greater income inequality. Lee (2026) finds stronger responses to monetary and fiscal policy at higher inequality, including larger state output responses to military spending.[2] For monetary policy, Ma (2023) instead finds larger output and employment responses to a monetary expansion in U.S. states with lower income

[1] A related literature examines how monetary and fiscal policy shocks affect the distribution of income and consumption (Coibion, Gorodnichenko, Kueng and Silvia, 2017, Anderson, Inoue and Rossi, 2016, Auerbach, Gorodnichenko and Murphy, 2026). We examine how output and employment responses to government spending vary with lagged income inequality.

[2] Lee (2026) uses annual spending changes and the lagged top 10% income share. Our baseline uses two-year spending changes and measures inequality relative to the annual cross-state median.

inequality and develops a model linking this pattern to aggregate labor supply elasticity.[3] For interest rates, Miranda-Pinto, Murphy, Walsh and Young (2023) find that government bond yields rise less or fall more after fiscal stimulus in OECD countries with greater income inequality. Their model emphasizes households that borrow to maintain a minimum consumption level and use additional income to repay debt. Our state-level IV design uses differences in exposure to national military spending changes to estimate relative state responses. Our estimates instead show smaller output multipliers in high-inequality states, while baseline employment multipliers do not differ significantly.

Our local estimates also build on the regional multiplier literature. Nakamura and Steinsson (2014) estimate an open-economy relative multiplier using military spending, and Dupor and Guerrero (2017) extend that design with updated data. Suárez Serrato and Wingender (2016) use variation in federal spending induced by Census population revisions, and Chodorow-Reich (2019) surveys cross-sectional fiscal estimates. Bellifemine, Couturier and Jamilov (2025) develop a regional Keynesian cross linking regional MPCs and non-tradable employment shares to aggregate monetary transmission. We show that local output multipliers vary with top income concentration and examine the associated production and income responses.

Section 2 presents the data, empirical strategy, and main estimates. Section 3 examines output per worker, income responses, and state characteristics, and assesses robustness to alternative inequality measures and samples. Sections 4–6 develop and calibrate the model and use it to study income incidence. Section 7 concludes.

2 Empirical Analysis

In this section, we examine how local output and employment multipliers vary with income inequality. We first describe the state-level data and the empirical strategy based on national military spending changes. We then present the two-year estimates of government spending multipliers and use local projections to trace the responses at longer horizons.

2.1. Data

We combine state-level military spending with income shares, output, and employment to compare spending responses across the inequality distribution. The baseline panel covers all 50 states and the District of Columbia over 1966–2006. Appendix A provides detailed data sources and

[3] Cravino, Lan and Levchenko (2020) examine differences in price stickiness across households' consumption baskets. In their model, greater income inequality strengthens monetary transmission, although the effect is modest for realistic changes in inequality.

construction.

Military Spending. Our primary measure of government spending is military procurement from the DD-350 Individual Contract Action Report database, as compiled by Nakamura and Steinsson (2014). These forms document military purchases exceeding $10,000 (pre-1983) and $25,000 (post-1983), accounting for approximately 90% of total military purchases. The DD-350 forms identify prime contractors and the location where most of the work was performed.

We use the sum of total military procurement and compensation of military employees by state and year for 1966–2006 as the baseline spending measure, following Nakamura and Steinsson (2014). Military procurement is particularly suitable for studying fiscal multipliers because it is driven by national security considerations—primarily the Vietnam War buildup (1966–1968), the post-Vietnam drawdown (1969–1975), the Carter-Reagan buildup (1979–1986), and the post-Cold War drawdown (1987–1997)—rather than by local economic conditions.[4]

Income Inequality. Our main inequality measure is the top 10% income share from the Frank-Sommeiller-Price series (Frank et al., 2015), derived from IRS *Statistics of Income* (SOI) tax data. This comprehensive panel covers all 50 states and the District of Columbia from 1917 onward. The pre-tax adjusted gross income includes wages, salaries, capital income (dividends, interest, capital gains), and entrepreneurial income. Following the world top incomes methodology of Piketty and Saez (2003), we use top decile shares as the primary measure and top 5% and top 1% shares as alternatives.

Output and Employment. State output is measured by GDP by state (Gross State Product) from the Bureau of Economic Analysis (BEA). State employment comes from the Bureau of Labor Statistics (BLS) Current Employment Statistics (CES) program, which provides monthly establishment-survey data aggregated to annual state totals. Population data are from the Census Bureau intercensal estimates. All monetary variables are deflated by the national CPI-U from BLS and expressed in real per capita terms.

Appendix Table A.1 reports summary statistics for the main variables and the additional measures used in the mechanism and robustness exercises. The top 10% income share has a mean of 37.1% with an overall standard deviation of 5.35 percentage points. The *within-state* standard deviation (4.59) exceeds the

textitbetween-state standard deviation (2.76), indicating that temporal variation within states—driven primarily by the secular increase in top income concentration since the early 1980s—is the

[4]For the extended sample through 2014, we use contract-only data from Dupor and Guerrero (2017), which provides consistent coverage of the post-2006 period.

dominant source of variation. The measure ranges from 20.4% in Alaska in 1976 to 59.4% in Florida in 2005, reflecting both the low inequality of the compressed postwar era and the high inequality of the more recent period.[5]

2.2. Empirical Strategy

To assess whether spending responses differ with inequality, we first estimate the output and employment responses without the inequality interaction. This provides a common benchmark before allowing the response to vary across the income distribution. Following Nakamura and Steinsson (2014), we use two-year changes:

$$\frac{Y_{it} - Y_{it-2}}{Y_{it-2}} = \alpha_i + \gamma_t + \beta \frac{G_{it} - G_{it-2}}{Y_{it-2}} + \varepsilon_{it}, \tag{2.1}$$

where Y_{it} is real per capita output in state i at time t, G_{it} is real per capita military spending, and α_i and γ_t are state and year fixed effects. For employment, the dependent variable is the proportional change in CES payroll employment divided by population, $(E_{it}/N_{it} - E_{it-2}/N_{it-2})/(E_{it-2}/N_{it-2})$, where E denotes payroll employment and N population. The spending denominator remains lagged real per capita output Y_{it-2}. The two-year change specification smooths annual noise and allows time for local demand propagation. An increase in spending equal to one percent of lagged output raises output, or the employment rate, by approximately β percent. The employment response is a percent change in the rate, not a percentage-point change.

The year fixed effects γ_t absorb effects common across states in each year, including the common effects of monetary policy, federal taxes, oil prices, and national business cycle conditions. The multiplier is estimated from *relative* variation across states within the same year. Heterogeneous state responses to common shocks need not be absorbed, so the identifying assumption still concerns the relationship between the instruments and relative state outcomes. State fixed effects α_i absorb time-invariant differences in the outcome growth rates.

Identification. Since military spending at the state level may respond to local economic conditions—through political allocation of contracts or endogenous base closures—we instrument for state-level spending using the interaction of national military spending changes with state-specific sensitivities. Formally, the instrument set consists of:

$$Z_{it} = \left\{ \mathbf{1}[i = s] \times \frac{G_t^{\text{nat}} - G_{t-2}^{\text{nat}}}{Y_{t-2}^{\text{nat}}} \right\}_{s=2}^{N},$$

[5] Appendix Figures A.1 and A.2 display the full time-series variation by state.

where G_t^{nat} and Y_t^{nat} are national military spending and output in real per capita terms, constructed from the state aggregates. The state-specific sensitivity arises because each state has a different historical specialization in defense production. States such as California, Connecticut, and Virginia receive disproportionately large shares of procurement during national buildups, while states with little defense industry receive very little.

The identifying assumption is that the United States does not embark on military buildups—such as those associated with the Vietnam War, the Soviet invasion of Afghanistan, or the post-September 11 conflicts—because states that receive a disproportionate share of procurement happen to be experiencing relative economic downturns. This is a weaker assumption than exogeneity of aggregate military spending, since time fixed effects absorb any correlation between national military buildups and the national business cycle. As Nakamura and Steinsson (2014) discuss, the key threat to identification would be if the political allocation of contracts were systematically directed toward economically distressed states within a given year. The procurement patterns are, however, largely determined by the location of existing defense contractors and military installations, which change slowly over time.

Inequality Interaction. The baseline specification imposes a common spending response across states. To allow that response to vary with income inequality, we center the inequality measure around its cross-state median in each year:

$$I_{it} = \text{Top10}_{it} - \text{median}_j[\text{Top10}_{jt}].$$

This centering ensures that I_{it} captures relative inequality—how unequal state i is compared with the typical state in year t—and the coefficient β_m in the interaction specification has a natural interpretation as the multiplier in the median-inequality state. We measure inequality at the beginning of the two-year spending window (I_{it-2}) to reduce contemporaneous feedback from spending to inequality.

Our main specification augments Equation (2.1) with an inequality interaction:

$$\frac{Y_{it} - Y_{it-2}}{Y_{it-2}} = \alpha_i + \gamma_t + \beta_m \frac{G_{it} - G_{it-2}}{Y_{it-2}} + \beta_d \, I_{it-2} \frac{G_{it} - G_{it-2}}{Y_{it-2}} + \delta \, I_{it-2} + \varepsilon_{it}, \tag{2.2}$$

where I_{it-2} is the deviation of the state's top 10% income share from the cross-state median at $t-2$. The coefficient β_m measures the multiplier at the median-inequality state. β_d captures how the multiplier varies with inequality. Both the spending variable and its interaction with inequality are treated as endogenous. The instrument set is expanded to include the interaction of the state-specific national spending instrument with lagged inequality: $\{Z_{it}\}$ and $\{Z_{it} \times I_{it-2}\}$. We treat lagged

Table 1: Baseline Results: Output and Employment Multipliers

	Output		Employment	
	(1)	(2)	(3)	(4)
$\Delta G_{it}/Y_{it-2}$	1.62	1.65^{**}	1.39^{*}	1.04^{**}
	(1.04)	(0.64)	(0.70)	(0.45)
$I_{it-2} \times \Delta G_{it}/Y_{it-2}$		-0.189^{***}		0.042
		(0.066)		(0.070)
Observations	1,989	1,989	1,989	1,989

Notes: IV estimates are from Equations (2.1) and (2.2). The inequality measure I_{it-2} is the deviation of the top 10% income share from the cross-state median at $t-2$. Instruments are state-specific national military spending (and its interaction with inequality). All regressions include state and year fixed effects. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years. $^{***}p<0.01$, $^{**}p<0.05$, $^{*}p<0.1$.

inequality I_{it-2} as exogenous conditional on the state and year fixed effects.

The implied multiplier for a state with inequality level I_{it-2} is $\mathscr{M}_i = \beta_m + \beta_d I_{it-2}$. If $\beta_d < 0$, states with above-median inequality have lower multipliers than states with below-median inequality. Section 3 asks through which channel.

Local Projection IV. To trace out the dynamic response of output and employment to spending shocks, and to examine how the inequality interaction evolves over time, we estimate local projection impulse responses (Jordà, 2005) at horizons $h = 0, 1, \ldots, 8$ years:

$$\frac{Y_{it+h} - Y_{it-2}}{Y_{it-2}} = \alpha_i^h + \gamma_t^h + \beta_m^h \frac{G_{it} - G_{it-2}}{Y_{it-2}} + \beta_d^h I_{it-2} \frac{G_{it} - G_{it-2}}{Y_{it-2}} + \delta^h I_{it-2} + \varepsilon_{it}^h, \tag{2.3}$$

estimated separately for each h, with the same instruments as in Equation (2.2). For employment, the left-hand side is the proportional change in the employment rate, while the spending denominator remains lagged output. The sequences $\{\beta_m^h\}$ and $\{\beta_d^h\}$ measure outcome responses relative to the *initial* two-year spending change. They are not ratios of cumulative output responses to cumulative spending responses. The horizon $h = 0$ reproduces the static two-year-change specification. The local projection approach accommodates the inequality interaction without imposing a VAR's dynamic structure (Cloyne et al., 2023).

2.3. Inequality and the Fiscal Multiplier

We first estimate the spending response without allowing it to differ with inequality. Columns (1) and (3) of Table 1 replicate the Nakamura and Steinsson (2014) specification. We use Driscoll–Kraay

standard errors for the panel IV regressions.[6] The output multiplier is 1.62 (s.e. 1.04, $p = 0.128$) and the employment multiplier is 1.39 (s.e. 0.70, $p = 0.055$). These point estimates are consistent with the original estimates in Nakamura and Steinsson (2014), but their Driscoll–Kraay standard errors exceed the corresponding state-clustered standard errors. The output point estimate implies that a one-dollar increase in real military spending per capita raises real local output per capita by approximately \$1.62. The employment point estimate implies that a spending increase equal to one percent of lagged output raises the employment rate by approximately 1.39 percent.

Columns (2) and (4) introduce the inequality interaction. The output multiplier at the median-inequality state is $\hat{\beta}_m = 1.65$ (s.e. 0.64), virtually unchanged from the unconditional estimate. The inequality interaction is $\hat{\beta}_d = -0.189$ (s.e. 0.066, $p = 0.007$). In economic terms, this means that a state whose top 10% income share is one percentage point above the cross-state median has an output multiplier that is 0.19 lower than in a state at the median. For illustration, a state whose top 10% income share is 5 percentage points higher than another state's has an implied output multiplier nearly one full point lower ($5 \times 0.19 \approx 0.95$). This is an economically large effect—it implies that the returns to fiscal stimulus are different across states with different income distributions, even within the same monetary union and under the same monetary and federal tax policy.

For employment, $\hat{\beta}_m = 1.04$ (s.e. 0.45) and $\hat{\beta}_d = 0.042$ (s.e. 0.070). The interaction is positive but statistically insignificant. Its approximate 95% confidence interval is $[-0.100, 0.185]$. The interval allows for both smaller and larger employment responses in high-inequality states. The contrast with the negative output interaction concerns the initial two-year window.

2.4. Dynamic Effects

We use local projections to examine how long the smaller output response in high-inequality states persists and whether employment responses differ at longer horizons. The coefficients measure the outcome responses relative to the initial two-year spending change.

Figure 1 presents the local projection IV estimates.[7] Panel (a) traces the output response coefficient over horizons $h = 0$ through $h = 8$. The response builds from 1.62 at $h = 0$ (the initial two-year window) to approximately 7.5 at horizons 4–5 before declining. Each coefficient uses the

[6] These standard errors allow for heteroskedasticity, serial correlation, and dependence across states (Driscoll and Kraay, 1998). We use Bartlett weights with maximum lag $L = \lfloor 4(T/100)^{2/9} \rfloor$, where T is the number of years in the estimation sample. This gives three lags in the baseline sample. Tests and confidence intervals use a t_{T-1} reference distribution. Appendix B.1 reports the corresponding state-clustered results and sensitivity to the lag choice. Driscoll–Kraay inference relies on a long time dimension and may be less reliable in the shorter samples.

[7] For local projections, we use $\max\{L, h+1\}$ lags in the Driscoll–Kraay covariance to allow for overlapping outcome changes, with L defined in footnote 6. The dynamic figures report pointwise 90% confidence bands, with 68% inner bands.

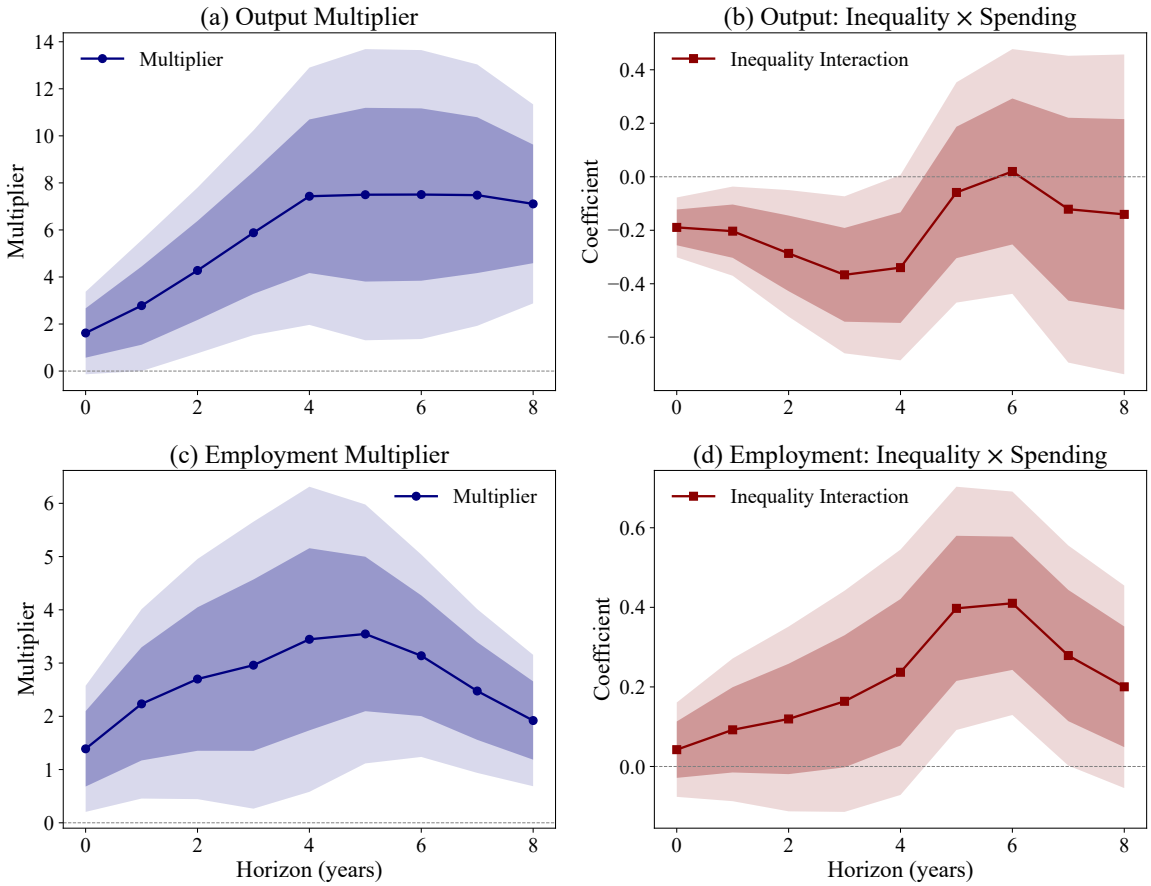


Figure 1: Local Projection Impulse Responses: Output and Employment

Notes: Local projection IV estimates from Equation (2.3) at horizons $h = 0, \ldots, 8$. Top row: output (percent change in real per capita output relative to $t-2$). Bottom row: employment (percent change in the employment rate relative to $t-2$). Left column: spending coefficient from the specification without the inequality interaction. Right column: inequality interaction $\hat{\beta}_d^h$ from Equation (2.3). Shaded areas represent 68% (inner, darker) and 90% (outer, lighter) pointwise confidence bands based on Driscoll–Kraay standard errors and t_{T-1} critical values. The maximum lag is $\max\{\lfloor 4(T/100)^{2/9} \rfloor, h+1\}$, where T is the number of years in the estimation sample.

initial spending change as its denominator. The peak is not a contemporaneous or cumulative dollar multiplier. The pattern indicates a persistent output response, although it does not by itself distinguish the channels of propagation.

Panel (b) of Figure 1 shows the inequality interaction $\hat{\beta}_d^h$, which measures how the output response varies with inequality. The interaction is negative and statistically significant at short horizons ($h = 0$ through $h = 3$), with the point estimate declining from -0.19 at $h = 0$ to -0.37 (s.e. 0.17) at $h = 3$. This means that the gap in output responses between low- and high-inequality states widens as the spending propagates through the local economy. From $h = 4$ onward, the 90% confidence bands include zero.

The bottom row of Figure 1 presents employment responses. The response coefficient peaks

at approximately 3.5 at horizons 4–5. The inequality interaction is not significantly different from zero through $h = 4$, but is positive and significant at the 10% level at $h = 5, 6, 7$. At $h = 5$ and $h = 6$, it is also significant at the 5% level.[8] The initial output-employment contrast therefore does not imply equal employment responses at every horizon. Section 3.1 examines the corresponding output-per-worker responses.

3 Mechanisms and Robustness

In this section, we examine the production and income responses behind the smaller output multiplier in high-inequality states and assess the robustness of the inequality interaction. We first decompose the output response into employment and output per worker, then examine personal income and its components. We next consider observed state characteristics and alternative inequality measures, spending measures, and sample periods.

3.1. Output per Worker and Employment

An output response to government spending can differ because the employment response differs or because output per worker differs. We estimate both responses on the same sample, so that the decomposition accounts for the output interaction without imposing a zero employment interaction.

Decomposing the Output Response. Let $Q_{it} = N_{it} Y_{it}$ denote real gross state product, where Y_{it} is real per capita output, E_{it} is payroll employment, and N_{it} is state population. Output per capita factors exactly into labor productivity and the employment rate,

$$\Delta \log Y_{it} = \underbrace{\Delta \log \frac{Q_{it}}{E_{it}}}_{\text{output per worker}} + \underbrace{\Delta \log \frac{E_{it}}{N_{it}}}_{\text{employment rate}}, \tag{3.1}$$

where Δ is the two-year difference used throughout. We estimate Equation (3.1) term by term using the specification of Section 2.2. Each log change is regressed on $\Delta G_{it}/Y_{it-2}$ and its interaction with the lagged top-10% share deviation, instrumented by the state-specific national spending interactions, with state and year fixed effects and Driscoll–Kraay standard errors. Because the three dependent variables sum exactly, are estimated on a common sample, and use identical instruments, the estimated coefficients—both the level and the interaction terms—add up exactly.

[8] The longer-horizon employment interactions depend on the lag choice. With three lags at every horizon, the p-values at $h = 5, 6, 7$ are 0.068, 0.056, and 0.186, compared with 0.035, 0.019, and 0.097 under the horizon-dependent rule. Thus the interactions at $h = 5$ and $h = 6$ remain significant at 10%, but not at 5%, with three lags. Appendix B.1 reports the lag comparisons and state-clustered results.

Table 2: The Productivity Decomposition of the Output Multiplier

Dependent variable	$\hat{\beta}_m$	s.e.	$\hat{\beta}_d$	s.e.
Δ log output per capita	1.590	(0.624)	-0.167^{**}	(0.066)
Δ log output per worker	0.543	(0.310)	-0.206^{**}	(0.077)
Δ log employment rate	1.047	(0.454)	0.039	(0.070)

Notes: IV estimates, 1968–2006, $N = 1{,}989$. The three rows are the terms of the identity $\Delta \log Y = \Delta \log(Q/E) + \Delta \log(E/N)$, estimated separately on a common sample with identical instruments, so the coefficients add up exactly. $\hat{\beta}_m$ is the response at median inequality. $\hat{\beta}_d$ is the interaction with the lagged top-10% share deviation. State and year fixed effects are included and Driscoll–Kraay standard errors are reported. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Two-Year Responses. Table 2 reports the decomposition. The employment-rate interaction is 0.039 (s.e. 0.070), while the output-per-worker interaction is −0.206 (s.e. 0.077).[9] Their sum is the log output interaction, −0.167 (s.e. 0.066). The negative output interaction is therefore accounted for by output per worker, partly offset by the positive employment point estimate. Both the output and output-per-worker interactions are significant at the 5% level.[10] These log-change coefficients differ from the proportional-change specification in Table 1.

The magnitudes are economically large. Across the 10th and 90th percentiles of the top-10% share deviation (a spread of 7.6 percentage points), a spending increase worth one percent of output raises output per worker by roughly 1.2 percent in the least unequal states and has a point estimate of −0.35 percent in the most unequal states. Employment responses are not significantly different in these comparisons.

Responses at Longer Horizons. The two-year decomposition leaves open whether output per worker continues to account for the smaller output response in high-inequality states as employment adjusts. Figure 2 reports local-projection impulse responses for each term of equation Equation (3.1).[11]

The difference in output-per-worker responses is present in the initial two-year window and then grows. The interaction is −0.21 (s.e. 0.08) at $h = 0$, reaches −0.50 (s.e. 0.18) at $h = 4$, and decays thereafter, remaining significant through $h = 4$. The 90% bands include zero from $h = 5$ onward. The employment-rate interaction is insignificant through $h = 4$, then positive and significant at

[9] Payroll employment is measured by place of work, consistent with gross state product. Using household employment measured by place of residence gives an output-per-worker interaction of −0.193 (s.e. 0.099), significant at the 10% level. Comparable state hours data are available only for manufacturing. Appendix B.2 examines average weekly manufacturing hours.

[10] The output-per-worker interaction remains negative and significant after adding spending interactions with college attainment, high-school attainment, and the construction share. Appendix B.2 reports these estimates.

[11] We use outcomes in logs so that the identity—and hence the adding-up of coefficients—holds at every horizon. The outcome at horizon h is $X_{i,t+h} - X_{i,t-2}$, so $h = 0$ reproduces Table 2.

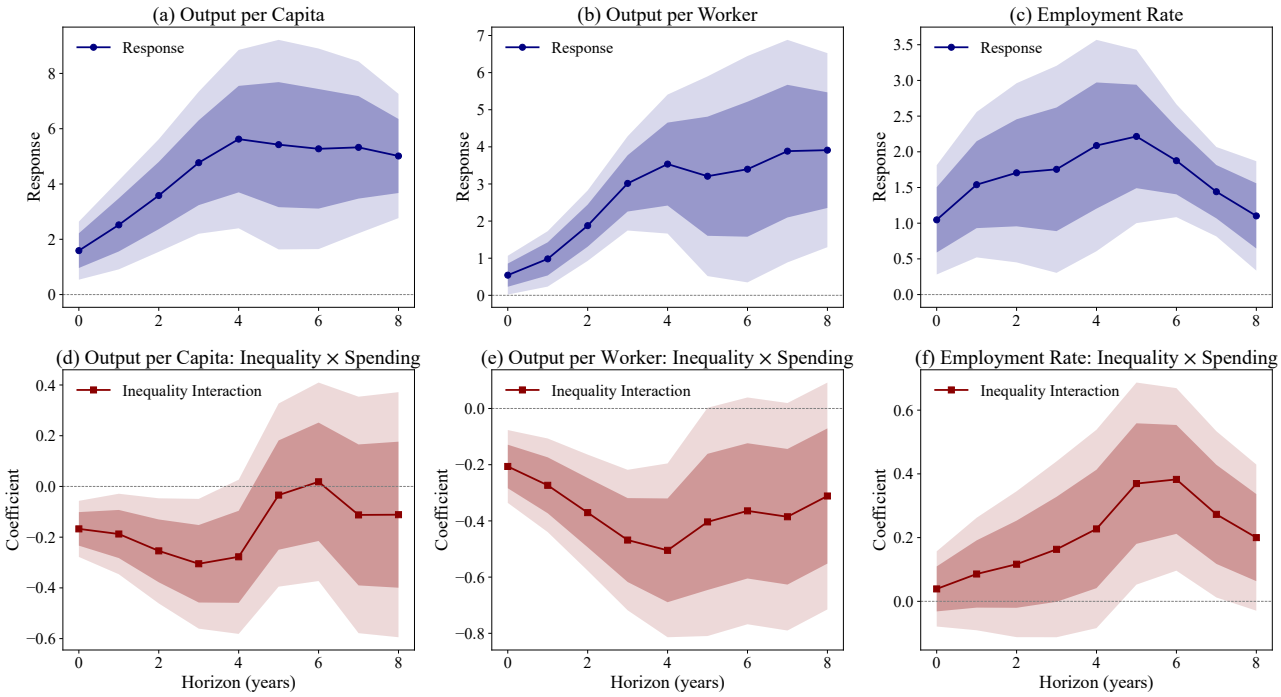


Figure 2: Local Projection Impulse Responses: The Productivity Decomposition

Notes: Local projection IV estimates at horizons $h = 0,\ldots,8$ for each term of the identity $\Delta \log Y = \Delta \log(Q/E) + \Delta \log(E/N)$. The outcome at horizon h is $X_{i,t+h} - X_{i,t-2}$, so $h = 0$ corresponds to the two-year change reported in Table 2. Panels (a)–(c) report the response to a spending increase worth one percent of output, $\hat{\beta}_m^h$. Panels (d)–(f) report the inequality interaction, $\hat{\beta}_d^h$. Shaded areas represent 68% (inner, darker) and 90% (outer, lighter) pointwise confidence bands based on Driscoll–Kraay standard errors.

the 10% level at $h = 5, 6, 7$. The timing leaves room for both an early change in worker composition and later adjustment in productive inputs, but does not identify their separate contributions. All horizons use the initial spending change as the denominator.

3.2. Income Responses to Military Spending

We next examine the income gains associated with the smaller output response in high-inequality states. Additional income can support consumption and further production. We use the state income accounts to measure how personal income and its components respond to military spending across the inequality distribution.

The BEA state personal income accounts measure income by source. Personal income combines earnings by place of work (wages and salaries, supplements, and proprietors' income), dividends, interest, rent, and transfer receipts, less social-insurance contributions and plus the residence adjustment.[12] We estimate Equation (2.2) using each component's two-year change in real per

[12] Personal income is recorded by residence and excludes corporate retained earnings. Earnings by place of work enter personal income together with contributions and a residence adjustment. Proprietors' income mixes returns to labor and capital. Appendix A.4 documents the income components and their construction.

Table 3: Personal Income Responses to Military Spending

	IV estimates		Implied response	
Dependent variable: two-year change in	$\hat{\beta}_m$	$\hat{\beta}_d$	5 pp below	5 pp above
Personal income	1.08^{***}	-0.108^{**}	1.62	0.54
	(0.36)	(0.048)	(0.40)	(0.46)
Earnings by place of work	1.18^{***}	-0.100^{*}	1.67	0.68
	(0.35)	(0.051)	(0.45)	(0.41)
Dividends, interest, and rent	0.13	0.002	0.12	0.15
	(0.08)	(0.016)	(0.11)	(0.12)
Transfer receipts	−0.04	−0.005	−0.01	−0.06
	(0.02)	(0.007)	(0.04)	(0.05)
Less: contributions for social insurance	0.11^{**}	0.001	0.10	0.11
	(0.05)	(0.007)	(0.05)	(0.07)
Residence adjustment	-0.09^{**}	−0.005	−0.07	−0.12
	(0.04)	(0.018)	(0.11)	(0.08)

Notes: IV estimates of Equation (2.2), 1968–2006, $N = 1{,}989$. The dependent variable is the two-year change in the real per capita income component divided by lagged real per capita output, so $\hat{\beta}_m$ is dollars of that income per dollar of military spending in the median-inequality state and $\hat{\beta}_d$ is the change in that response per percentage point of the lagged top 10% share deviation. The last two columns report $\hat{\beta}_m \mp 5\hat{\beta}_d$, the implied responses in states five percentage points below and above the median, with standard errors computed from the estimated covariance matrix of $(\hat{\beta}_m, \hat{\beta}_d)$. The BEA state personal income accounts record income by residence. Earnings by place of work include wages and salaries, supplements to wages and salaries, and proprietors' income. They are adjusted for social-insurance contributions and place of residence as shown in the table. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

capita dollars divided by lagged real per capita output. Thus $\hat{\beta}_m$ is dollars of income per dollar of military spending and $\hat{\beta}_d$ is the change in that response per percentage point of the lagged top 10% share deviation. Components are estimated on common samples with identical regressors and instruments, so their coefficients add to the total up to numerical precision.

Table 3 reports the income responses. In the median state, a dollar of military spending generates about \$1.08 of resident personal income. The implied response is \$1.62 in a state five percentage points below the median top 10% share and \$0.54 in a state five points above. The personal-income interaction is −0.108 (s.e. 0.048), significant at the 5% level.

The difference in the personal-income point estimates mainly reflects earnings. Earnings by place of work respond by \$1.67 per spending dollar in the low-inequality comparison and \$0.68 in the high-inequality comparison. The earnings interaction is −0.100 (s.e. 0.051, $p = 0.057$), providing suggestive evidence of a smaller earnings gain. Dividends, interest, and rent respond by \$0.12 and \$0.15, with a statistically insignificant interaction of 0.002 (s.e. 0.016).[13]

[13] Appendix B.1 reports the inequality interactions for the three earnings components. Appendix B.3 reports income

Table 4: State Characteristics and the Inequality Interaction

Characteristic Z	Type	Baseline $\hat{\beta}_d$	With $Z \times \Delta G$ $\hat{\beta}_d$	Attenuation	Coefficient on $Z \times \Delta G$
Earnings share	Income share	−0.189	−0.115	39%	6.37
		(0.066)	(0.086)		(10.68)
Union density	Labor market	−0.189	−0.137	27%	0.070
		(0.066)	(0.074)		(0.041)
CPS income ratio	Income distribution	−0.189	−0.194	-3%	4.41
		(0.066)	(0.056)		(4.17)
Property income ratio	Income share	−0.189	−0.215	-13%	−6.63
		(0.066)	(0.080)		(5.91)

Notes: Each row uses the same sample in the baseline and augmented specifications ($N = 1,989$). The added state characteristic and its spending interaction enter with corresponding instruments. Attenuation is $1 - |\hat{\beta}_d^{aug}|/|\hat{\beta}_d^{base}|$, computed before rounding and allowed to be negative. Earnings share is earnings divided by earnings plus dividends, interest, and rent. The property-income ratio adds proprietors' income to dividends, interest, and rent, using the same denominator. The CPS income ratio is $p10/p50$, not an observed HtM share. Variables are lagged two years and centered on year medians. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

The estimates show smaller personal-income gains in high-inequality states and suggest smaller earnings gains. Their effect on consumption depends on which households receive the gains and those households' MPCs. Earnings accrue to households with different MPCs, so the distribution of these gains across household types remains unresolved by the income accounts.

3.3. State Characteristics

The smaller income gains in high-inequality states suggest examining state characteristics that may shape how spending generates income and supports consumption. Income shares describe the relative importance of earnings and property income, union density may affect wage responses, and the bottom-to-median income ratio describes incomes near the bottom relative to the median. We first allow spending responses to vary with these characteristics, then examine the inequality interaction after adding state characteristics as controls.

3.3.1. Spending interactions. Table 4 reports these comparisons. Each regression augments the baseline with an additional spending interaction $\Delta G \times Z$, the characteristic itself, and corresponding instruments. These coefficient changes are descriptive comparisons. They do not identify the fraction of a causal effect transmitted through each characteristic.

Earnings share. The earnings share is earnings by place of work divided by earnings plus

responses at longer horizons.

dividends, interest, and rent. Its spending interaction changes the inequality coefficient from −0.189 to −0.115, an attenuation of 39%. The additional interaction is positive (6.37) but imprecise (s.e. 10.68). The measure is an income share, not compensation/GDP or the marginal earnings response to spending. Appendix B.4 reports the corresponding GDP-denominated measures: compensation/GDP changes the inequality coefficient to −0.204, while earnings/GDP changes it to −0.174. Thus the 39% attenuation does not generalize to these measures and does not identify the contribution of labor-income incidence.

Union density. Union density is the share of nonagricultural wage and salary employees who are union members, from UnionStats. Adding its spending interaction changes the inequality coefficient from −0.189 to −0.137, a 27% reduction. The union interaction is positive but imprecise (0.070, s.e. 0.041). Its p-value is just above 0.10. The coefficient change is descriptive and does not identify the causal contribution of labor institutions.

Bottom-to-median income ratio. Adding the spending interaction with the CPS 10th/50th income percentile ratio does not attenuate the inequality coefficient: it moves from −0.189 to −0.194, an attenuation of −3% after rounding. The additional interaction is positive but imprecise (4.41, s.e. 4.17). A higher ratio means that bottom incomes are closer to the median. It is not an observed increase in the HtM population share. The result does not exclude household composition as a mechanism. It shows that this income-distribution measure does not account for the top-share interaction.

Property and proprietors' income ratio. This measure divides dividends, interest, rent, and proprietors' income by earnings plus dividends, interest, and rent. Its inclusion changes the inequality coefficient from −0.189 to −0.215. Its spending interaction is negative (−6.63) but imprecise (s.e. 5.91). Since proprietors' income also appears within earnings and resident property income does not measure nonlocal flows, this ratio does not identify a leakage channel.

3.3.2. Additional controls. Next, we add state characteristics as lagged level controls in the baseline specification. These regressions assess the sensitivity of the inequality interaction to additional controls. The inequality coefficient remains close to the baseline when controlling separately for the earnings share, the property and proprietors' income ratio, union density, or house price levels. Appendix B.5 reports these estimates.

State tax progressivity may also be correlated with inequality. Ferriere and Navarro (2025) show that multipliers are larger when additional spending is financed with higher tax progressivity. We add lagged state tax progressivity, measured by the top state marginal income tax rate minus the average state wage tax rate, as a level control. TAXSIM coverage begins in 1979, leaving 1,326

Table 5: Robustness Checks

Specification	$\hat{\beta}_m$	s.e.	$\hat{\beta}_d$	s.e.	N
Baseline (top 10%)	1.65	(0.64)	−0.189	(0.066)	1,989
Alternative inequality measures					
Top 5% share	1.32	(0.62)	−0.15	(0.08)	1,989
Top 1% share	1.32	(0.65)	−0.19	(0.09)	1,989
Gini coefficient	1.41	(0.61)	−24.5	(23.7)	1,950
Wage p90/p10	1.41	(0.58)	−0.008	(0.060)	1,989
Additional control					
Tax progressivity	1.98	(0.88)	−0.12	(0.14)	1,326
Extended sample					
DG 1966–2014	1.45	(0.51)	−0.19	(0.05)	2,350

Notes: Each row reports IV estimates from a separate regression. $\hat{\beta}_m$ and $\hat{\beta}_d$ are the multiplier at median inequality and the inequality interaction, respectively. The baseline specification (first row) corresponds to column (2) of Table 1. For alternative inequality measures, the interaction uses the specified measure instead of the top 10% share. For the Gini, the coefficient is large because the Gini is measured in [0,1]. The tax progressivity specification adds the lagged difference between the top state marginal income tax rate and the average state wage tax rate as a level control. Wage p90/p10 uses the CPS weekly wage 90/10 ratio as the inequality measure. Extended sample uses contract-only spending from Dupor and Guerrero (2017). Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years.

observations. Table 5 reports an inequality interaction of −0.12 (s.e. 0.14), which is statistically insignificant. The tax progressivity coefficient is 0.007 (s.e. 0.005). Since tax progressivity enters in levels, this coefficient describes its conditional association with output growth. The specification leaves its effect on the spending response unresolved.

3.4. Alternative Inequality Measures and Samples

The state-characteristic comparisons retain the top 10% income share and the baseline military spending measure. We next examine how the inequality interaction changes with other inequality measures, the spending definition, and the sample period. Table 5 reports these comparisons alongside the tax-progressivity control discussed above.

Inequality measures. We first examine whether the result depends on using the top 10% income share, our preferred measure. Replacing it with the top 5% share yields $\hat{\beta}_d = -0.15$ (s.e. 0.08, $p = 0.058$). The top 1% share gives $\hat{\beta}_d = -0.19$ (s.e. 0.09, $p = 0.046$). The Gini interaction is negative but imprecise (−24.5, s.e. 23.73, $p = 0.309$). Its large magnitude reflects the 0–1 scale of the Gini. Employment interactions remain insignificant for the top-income-share measures, while the Gini employment interaction is negative and significant ($p = 0.049$). Thus the negative output relationship extends to other top income shares, but the evidence differs across inequality measures.

We next examine whether the result extends to general wage dispersion. The CPS wage 90/10

ratio yields an interaction of −0.008 (s.e. 0.06), which is not statistically significant. Thus this measure does not reproduce the top-share result. It does not establish that wage dispersion has no role, especially given the uncertainty and differences in the income concepts.

Spending measure and sample period. To examine whether the negative interaction extends beyond the baseline period, we use contract-only spending from Dupor and Guerrero (2017) over the extended 1966–2014 period. The output multiplier is 1.45, and the inequality interaction is −0.19 (s.e. 0.05, $p < 0.001$), qualitatively consistent with the baseline. Appendix B.6 presents local projection impulse responses for this extended sample. This exercise changes both the sample period and the spending definition.

To separate the spending definition from the sample period, Appendix B.7 holds the estimation sample fixed: using contracts only on the original 1968–2006 estimation sample yields an output multiplier of 1.59 (s.e. 0.55) and an inequality interaction of −0.259 (s.e. 0.094). The employment interaction is 0.010 (s.e. 0.063). The national spending instruments are reconstructed from the contracts-only measure. Thus, the negative output interaction is not explained by including military compensation in the baseline spending variable.

Aggregate evidence. Using military news shocks (Ramey, 2011, Ramey and Zubairy, 2018), Appendix C reports larger U.S. cumulative output multipliers in periods of lower income inequality. These estimates concern aggregate spending responses under national monetary and fiscal policies, which differ from the relative state responses estimated here.

3.5. Summary

The evidence associates the smaller output multiplier in high-inequality states with a smaller output-per-worker response and smaller personal-income gains. The state-characteristic regressions leave the structural source of these differences unresolved. The income accounts describe income sources, while the consumption response depends on the disposable income reaching households with different MPCs. In Section 4, we use a model to examine how production structure and labor-market frictions affect the income generated by spending and its distribution across households.

4 Model

Our empirical analysis documents smaller output multipliers in states with greater income concentration, while baseline employment multipliers do not differ significantly. The smaller output response is accounted for by output per worker, and personal-income gains, driven by earnings, are smaller in high-inequality states. The income accounts describe where income comes from,

whereas the consumption response depends on the disposable income reaching households with different marginal propensities to consume (MPCs). We use a two-agent New Keynesian model to examine whether the distribution of spending-induced income can account for this pattern.

In the model, the disposable income reaching HtM households per dollar of spending depends on both the total household disposable-income gain and the share received by these households.[14] We hold their population share fixed and examine how production structure and labor-market frictions affect this income.

To determine income incidence, the model combines capital–skill complementarity with segmented search-and-matching labor markets.[15] The production structure determines how the additional factor demand generated by spending, and hence the income it creates, is distributed across low-skill workers, high-skill workers, and capital owners. Search frictions and wage rigidities govern hiring in each labor-market segment, while the composition of employment, hours, and capital utilization determine how much each hire adds to output. These margins matter because employment multipliers do not differ significantly with inequality in the baseline estimates: the model should be able to lower the output response without mechanically lowering headcount employment, as in the empirical productivity decomposition (Section 3.1). The remainder of this section describes the model. Section 5 calibrates its steady state, and Section 6 uses it to study the incidence channel and its link to income concentration.

4.1. Households

The economy is populated by a unit mass of households. A fraction λ are hand-to-mouth households, denoted by H, and a fraction $1-\lambda$ are Ricardian households, denoted by R. Hand-to-mouth households supply low-skill labor and consume their current labor income. Ricardian households supply high-skill labor and save in government bonds and installed capital. Following Merz (1995) and Andolfatto (1996), each type is a large family whose members perfectly insure one another against idiosyncratic employment risk. At date t, a fraction $u_{i,t}$ of type-i members is unemployed and the remaining fraction $1-u_{i,t}$ is employed. An employed member supplies $n_{i,t}$ hours and e_i efficiency units of labor per hour.

[14] Appendix D presents the household-income decomposition and distinguishes the HtM population share from its share of additional income.

[15] Dolado, Motyovszki and Pappa (2021) combine capital–skill complementarity and asymmetric search-and-matching frictions to study the effects of monetary policy on labor income inequality. We use these features to examine how fiscal multipliers vary with income concentration.

Family $i \in \{H, R\}$ has preferences

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \frac{X_{i,t}^{1-\sigma} - 1}{1-\sigma}, \qquad X_{i,t} = C_{i,t} - \psi_i (1 - u_{i,t}) e_i \frac{n_{i,t}^{1+\xi_i}}{1+\xi_i},$$

where $C_{i,t}$ is family consumption per member. These Greenwood–Hercowitz–Huffman preferences (Greenwood, Hercowitz and Huffman, 1988) eliminate wealth effects on hours, so that hours depend only on the after-tax wage. The intratemporal optimality condition is

$$\psi_i n_{i,t}^{\xi_i} = (1 - \tau_t) w_{i,t}, \qquad i \in \{H, R\}, \tag{4.1}$$

where $w_{i,t}$ is the real wage per efficiency unit and τ_t is the proportional tax rate on labor and profit income.

4.1.1. Hand-to-mouth households. The hand-to-mouth family does not save and owns no claims on firms, so its consumption equals its disposable labor income:

$$C_{H,t} = (1 - \tau_t)(1 - u_{H,t}) e_H w_{H,t} n_{H,t}. \tag{4.2}$$

Changes in current disposable labor income therefore pass one-for-one into this family's consumption, which gives it a marginal propensity to consume of one.

4.1.2. Ricardian households. The Ricardian family smooths consumption by investing in one-period nominal government bonds and in installed capital. Let $b_{R,t}$ denote the real quantity of bonds purchased per Ricardian household at date t. One unit invested at t pays the gross nominal return R_t at $t+1$, so the realized real return is R_t / Π_{t+1}. Let $k_{R,t}$ denote the family's end-of-period holding of installed capital per member and Q_t its relative price. The real budget constraint is

$$\begin{aligned} C_{R,t} + b_{R,t} + Q_t k_{R,t} = {} & (1 - \tau_t)(1 - u_{R,t}) e_R w_{R,t} n_{R,t} + \frac{R_{t-1}}{\Pi_t} b_{R,t-1} \\ & + \left[Q_t + r_t^k \nu_t - \delta(\nu_t) \right] k_{R,t-1} + \frac{D_t}{1-\lambda}, \end{aligned} \tag{4.3}$$

where r_t^k is the rental rate per unit of capital services, ν_t is utilization, D_t denotes aggregate after-tax dividends, and $\delta(\nu_t) = \delta_0 + \delta_1 \nu_t^{\delta_2}$ is utilization-dependent depreciation. The payoff on installed capital in Equation (4.3) thus includes its resale value and rental income, net of depreciation. The Ricardian family chooses consumption, bond and capital holdings, and utilization to maximize utility subject to this constraint. The optimality conditions, including the stochastic discount factor and the no-arbitrage condition between bonds and capital, are in Appendix E.

4.1.3. Ownership concentration and measured income. To construct a model measure of top-income concentration, we separate the ownership of financial claims from the sharing of consump-

tion risk. Let $\chi \leq 1-\lambda$ denote the population share of Ricardian members who hold the family's financial claims. These owner-members receive the associated financial and profit income in the cross-sectional income distribution, but all resources are pooled within the Ricardian family, as in Equation (4.3), so all Ricardian members consume the same amount $C_{R,t}$.

Define per-capita disposable labor income for the two families as

$$y_{H,t} = (1-\tau_t)(1-u_{H,t})e_H w_{H,t} n_{H,t}, \qquad y^L_{R,t} = (1-\tau_t)(1-u_{R,t})e_R w_{R,t} n_{R,t}. \tag{4.4}$$

Aggregate financial and profit income accruing to the Ricardian family, excluding valuation changes in installed capital, is

$$\mathscr{F}_t = \left(\frac{R_{t-1}}{\Pi_t} - 1\right) B_{g,t-1} + \left[r^k_t \nu_t - \delta(\nu_t)\right] K_t + D_t. \tag{4.5}$$

The cross-sectional income distribution therefore consists of three groups with population masses and per-capita incomes

$$\left(\lambda,\, 1-\lambda-\chi,\, \chi\right), \qquad \left(y_{H,t},\, y^L_{R,t},\, y^L_{R,t} + \frac{\mathscr{F}_t}{\chi}\right).$$

In the quantitative analysis, χ equals the population share defining the top-income group. When owner-members are the highest-income group, their income share is

$$s^{\text{top}}_t = \frac{\chi y^L_{R,t} + \mathscr{F}_t}{\lambda y_{H,t} + (1-\lambda) y^L_{R,t} + \mathscr{F}_t}. \tag{4.6}$$

The income Gini coefficient is computed from the same three-point distribution. This ownership structure concentrates dividends and other financial income at the top while preserving complete consumption sharing within the Ricardian family.

4.2. Segmented labor markets

Labor markets are segmented by worker type, and each segment has its own search-and-matching frictions. At the beginning of period t, the measure of type-i job seekers is

$$S_{i,t} = u_{i,t-1} + \lambda_i (1 - u_{i,t-1}),$$

where λ_i is the exogenous job-separation probability. Matches are formed according to the constant-returns matching technology of Den Haan, Ramey and Watson (2000),

$$M_{i,t} = \frac{S_{i,t} V_{i,t}}{\left(S^{\alpha_i}_{i,t} + V^{\alpha_i}_{i,t}\right)^{1/\alpha_i}}, \tag{4.7}$$

$$f_{i,t} = \frac{M_{i,t}}{S_{i,t}}, \qquad q^f_{i,t} = \frac{M_{i,t}}{V_{i,t}},$$

where $V_{i,t}$ denotes vacancies, $f_{i,t}$ is the job-finding probability, and $q^f_{i,t}$ is the vacancy-filling probability. Unemployment evolves according to

$$u_{i,t} = u_{i,t-1} + \lambda_i(1 - u_{i,t-1}) - M_{i,t}.$$

Competitive labor agencies post vacancies and supply matched workers to intermediate-goods firms. The value of a filled type-i job is

$$J_{i,t} = (r_{i,t} - w_{i,t})e_i n_{i,t} + \beta(1-\lambda_i)\mathbb{E}_t J_{i,t+1},$$

where $r_{i,t}$ is the rental price paid by intermediate-goods firms per efficiency unit of labor, and free entry into vacancy posting implies

$$\iota_i = q^f_{i,t} J_{i,t},$$

where ι_i is the resource cost of posting a vacancy.

Real wages adjust gradually according to

$$\frac{w_{i,t}}{\bar{w}_i} = \left[\frac{w_{i,t-1}}{\bar{w}_i}\left(\frac{\bar{\Pi}}{\Pi_t}\right)^{d_i}\left(\frac{\Pi_{t-1}}{\Pi_t}\right)^{1-d_i}\right]^{\rho_{w,i}}\left(\frac{r_{i,t}}{\bar{r}_i}\right)^{(1-\rho_{w,i})\varepsilon_{w,i}},$$

where bars denote steady-state values. The parameter $\rho_{w,i}$ governs wage persistence, d_i determines the degree of indexation to steady-state rather than lagged inflation, and $\varepsilon_{w,i}$ governs the response of the wage to the value of the worker's marginal product.

4.3. Production and price setting

Production uses capital and the two types of labor in a nested CES technology that allows for capital–skill complementarity. Effective low- and high-skill labor inputs are

$$L_{H,t} = \lambda(1 - u_{H,t})e_H n_{H,t}, \qquad L_{R,t} = (1-\lambda)(1 - u_{R,t})e_R n_{R,t}. \tag{4.8}$$

Capital services are $\widetilde{K}_t = \nu_t K_t$. Intermediate-goods producers combine capital services and high-skill labor in an inner CES aggregate and then combine that aggregate with low-skill labor,

$$\begin{aligned} V_t &= \left[\omega_K \widetilde{K}_t^{\rho} + (1-\omega_K)L_{R,t}^{\rho}\right]^{1/\rho}, \\ Y_t &= \left[a L_{H,t}^{\zeta} + (1-a)V_t^{\zeta}\right]^{1/\zeta}. \end{aligned} \tag{4.9}$$

The elasticity of substitution between low-skill labor and the capital–high-skill composite is $1/(1-\zeta)$, while the elasticity between capital and high-skill labor is $1/(1-\rho)$. We assume capital–skill complementarity, $\zeta > 0 > \rho$, so that the outer nest is more substitutable than the inner nest.

Cost minimization yields

$$r_{H,t} = \mathrm{MC}_t Y_t^{1-\zeta} a L_{H,t}^{\zeta-1},$$
$$r_{R,t} = \mathrm{MC}_t Y_t^{1-\zeta} (1-a) V_t^{\zeta-\rho} (1-\omega_K) L_{R,t}^{\rho-1},$$
$$r_t^k = \mathrm{MC}_t Y_t^{1-\zeta} (1-a) V_t^{\zeta-\rho} \omega_K \tilde{K}_t^{\rho-1},$$

where MC_t denotes real marginal cost.

Intermediate-goods producers are monopolistically competitive with elasticity of substitution η across differentiated goods and set prices subject to Rotemberg adjustment costs (Rotemberg, 1982), which yield a standard Phillips curve (Appendix E).

4.4. Fiscal and monetary policy

Government purchases follow

$$g_t = \rho_G g_{t-1} + \varepsilon_t^G, \qquad g_t \equiv \log\left(\frac{G_t}{\bar{G}}\right), \tag{4.10}$$

where ε_t^G is an exogenous innovation. The fiscal authority issues one-period nominal debt and levies a common tax rate on labor and firm profits. Its flow budget constraint is

$$B_{g,t} = \frac{R_{t-1}}{\Pi_t} B_{g,t-1} + G_t - \tau_t \left(w_{H,t} L_{H,t} + w_{R,t} L_{R,t} + \mathscr{P}_t\right),$$

and the tax rate responds to the real value of outstanding debt,

$$\tau_t = (1-\rho_\tau)\bar{\tau} + \rho_\tau \tau_{t-1} + \phi_B \left[\frac{R_{t-1}}{\bar{R}} \frac{\bar{\Pi}}{\Pi_t} \frac{B_{g,t-1}}{\bar{B}_g} - 1\right].$$

The monetary authority sets the nominal interest rate according to

$$\log\left(\frac{R_t}{\bar{R}}\right) = \rho_R \log\left(\frac{R_{t-1}}{\bar{R}}\right) + (1-\rho_R)\left[\phi_\Pi \log\left(\frac{\Pi_t}{\bar{\Pi}}\right) + \phi_Y \log\left(\frac{Y_t}{\bar{Y}}\right)\right].$$

Capital production, profits and dividends, market clearing, the resource constraint, the definition of equilibrium, and the solution method are collected in Appendix E.

5 Calibration

We first calibrate the baseline steady state at a quarterly frequency. Table 6 reports the key parameters, and Appendix F documents all parameters, their sources, and the internally determined values. The baseline economy targets the median top 10% income share in our empirical sample. The parameters governing tax and monetary-policy responses and wage adjustment are chosen in Section 6.3 to match the empirical multiplier levels and inequality interactions, which separates the steady-state calibration from the moments used to determine fiscal responses.

Table 6: Key Calibrated Parameters

Parameter	Value	Description	Source or target
Panel A: Households and labor markets			
λ	0.75	Hand-to-mouth population share	College-attainment split
β	0.99	Quarterly discount factor	Smets and Wouters (2007)
ξ_H, ξ_R	1/3	Inverse intensive-margin Frisch elasticities	Chetty et al. (2011)
u_H, u_R	0.060, 0.025	Unemployment rates	Current Population Survey
λ_H, λ_R	0.055, 0.022	Quarterly separation probabilities	Wolcott (2021); Dolado et al. (2021)
Panel B: Production			
a	0.60	Low-skill weight in outer CES nest	Baseline calibration
ζ	0.401	Outer-nest CES parameter	Krusell et al. (2000)
ρ	−0.495	Inner-nest CES parameter	Krusell et al. (2000)
ω_K	0.35	Capital weight in inner CES nest	Dolado et al. (2021)
η	3	Elasticity across differentiated goods	De Loecker et al. (2020)
Panel C: Fiscal policy and distribution			
τ	0.25	Steady-state tax rate	Trabandt and Uhlig (2011)
$\bar{B}_g/(4\bar{Y})$	0.20	Debt relative to annual output	Baseline target
ρ_G, ρ_R	0.85, 0.85	Spending and policy-rate persistence	External calibration
χ	0.10	Population share assigned ownership income	Top-decile definition
s^{top}	0.36	Baseline top-10-percent income share	Sample median

Notes: Rates and probabilities are quarterly unless otherwise noted. Appendix Table F.1 reports all parameters, including those determined internally.

Households and labor markets. We set the share of hand-to-mouth, low-skill households to $\lambda = 0.75$, consistent with interpreting the two groups by college attainment near the beginning of our empirical sample, and the quarterly discount factor to $\beta = 0.99$. The inverse Frisch elasticities are $\xi_H = \xi_R = 1/3$. Normalizations of efficiency and hours imply a skill premium of 1.91, broadly consistent with existing empirical estimates. We target unemployment rates of $u_R = 0.025$ and $u_H = 0.060$, guided by education-specific unemployment in the CPS, and set quarterly separation probabilities at $\lambda_R = 0.022$ and $\lambda_H = 0.055$ (Wolcott, 2021, Cairó and Cajner, 2018).

Production. The capital–skill complementarity parameters $\zeta = 0.401$ and $\rho = -0.495$ are based on Krusell et al. (2000) and imply elasticities of substitution of 1.67 between low-skill labor and the capital–high-skill composite and 0.67 between capital and high-skill labor. The capital weight in the inner nest is $\omega_K = 0.35$, and the elasticity across differentiated goods $\eta = 3$ implies a gross markup of 1.5. In the baseline economy, we set the low-skill weight in the outer nest to $a = 0.60$, which yields an aggregate labor share of 0.61.

Government policies and income concentration. The steady-state tax rate is $\tau = 0.25$, government debt is 20 percent of annual output, and government purchases are about 20 percent of GDP.

Spending and the policy rate have persistence of 0.85. The owner group in the measured income distribution has population mass $\chi = 0.10$ (Section 4.1.3), and we choose the fixed operating cost Φ in the profit equation (Appendix E) so that the baseline top 10% income share is 36%, close to the sample median of 36.29%.[16]

6 Quantitative Experiments

The empirical analysis motivates income incidence as a possible explanation for the smaller output multiplier in high-inequality states. To assess this channel, we first use a controlled experiment to isolate how the allocation of additional income affects output and employment. We then change a structural parameter to study how income concentration and income incidence can move together.

Throughout, we compare fiscal multipliers. For year j, let $\overline{X}_j$ denote the average of the four quarterly levels and X^{ss} the corresponding steady-state level. We define

$$m_j^Y = \frac{\overline{Y}_j - Y^{ss}}{\overline{G}_j - G^{ss}}, \qquad m_j^N = \frac{\overline{N}_j - N^{ss}}{[\overline{G}_j - G^{ss}]/Y^{ss}}, \tag{6.1}$$

where N is the aggregate headcount employment rate. The year-2 ratios are the model counterparts of the two-year multipliers in our empirical estimates. In every comparison, the spending shock produces the same impact change in G/Y^{ss}, so that differences in multipliers reflect differences in propagation rather than in the size of the shock.

6.1. The income-incidence channel

The income generated by a spending expansion does not accrue evenly across households. The hand-to-mouth (low-skill) family H consumes its current disposable income, whereas the Ricardian (high-skill) family R can smooth consumption through saving. If a larger share of the additional income reaches the Ricardian family, private consumption rises by less, and so does the feedback from consumption to production and from production back to income.

A structural change in the economy would alter many margins at once, so we isolate this channel by changing only who receives incremental income. We keep the baseline steady state, production technology, labor-market structure, and policy rules. Let $W_{H,t} = \lambda y_{H,t}$, with $y_{H,t}$ defined in Equation (4.4), denote aggregate disposable income of the hand-to-mouth family. We introduce a zero-sum payment from H to R equal to a fraction of its deviation from steady state,

$$\mathcal{T}_t = \vartheta\left(W_{H,t} - \overline{W}_H\right), \qquad \vartheta \in [0,1]. \tag{6.2}$$

[16] The empirical target is a pre-tax income share including realized capital gains. The model measure uses disposable labor income, after-tax dividends, and capital income net of depreciation, as defined in Equation (4.6).

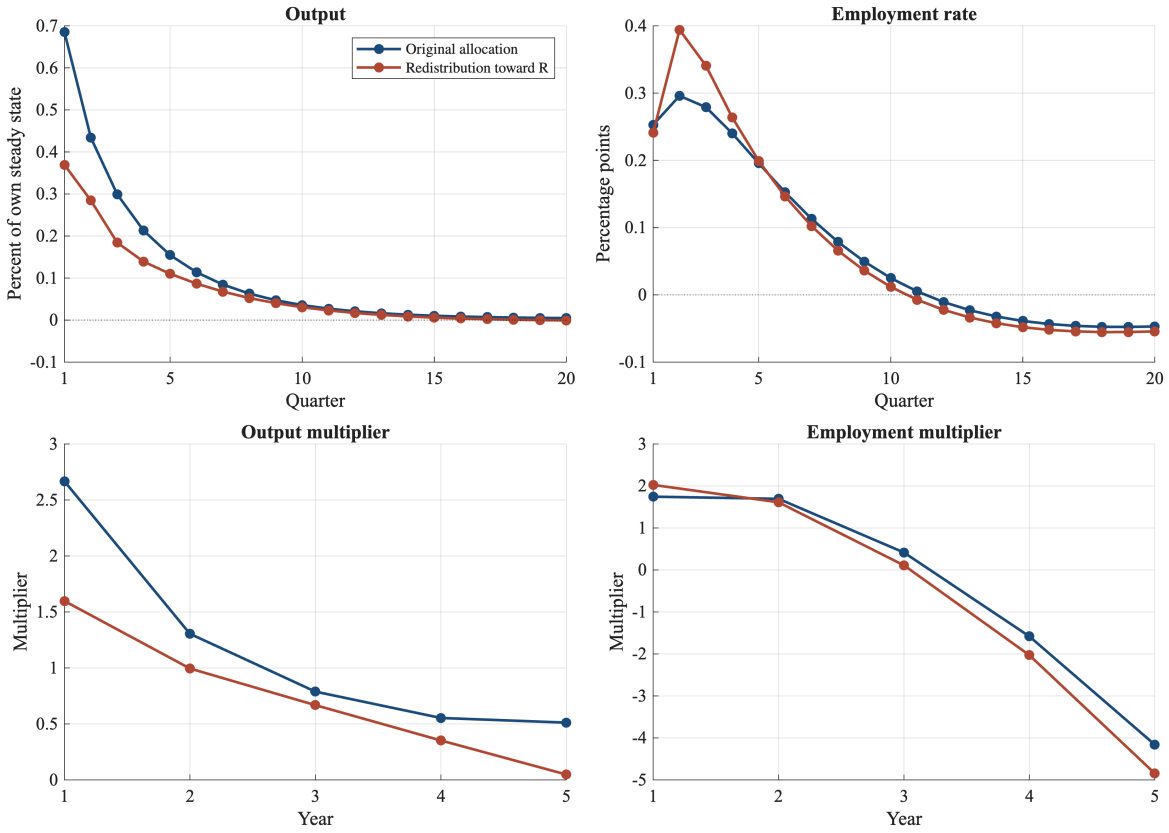


Figure 3: Fiscal Responses Under Alternative Allocations of Incremental Income

Notes: The original allocation corresponds to $\vartheta = 0$ and redistribution toward R to $\vartheta = 1$ in Equation (6.2). Both cases share the same steady state, production technology, labor-market conditions, policy parameters, and spending shock. Output responses are percentage deviations from steady state. Employment responses are percentage-point deviations. The bottom panels show the multipliers in Equation (6.1) for each year.

The payment enters the H budget Equation (4.2) as $-\mathcal{T}_t/\lambda$ and the R budget Equation (4.3) as $+\mathcal{T}_t/(1-\lambda)$. At $\vartheta = 0$, income is allocated as in the original economy. At $\vartheta = 1$, all of the additional wage income of the hand-to-mouth family goes to the Ricardian family, so that the former's disposable income stays at its steady-state level. Because $\mathcal{T}_t = 0$ in steady state, the two allocations share the same steady state and, in particular, the same income concentration. Because households take the payments as lump-sum, they do not distort marginal incentives: the hours condition Equation (4.1) and the production, wage-setting, and matching conditions are unchanged. Equilibrium wages, hiring, and hours can still adjust. Only the incidence of income differs across the two allocations, not production or recruitment technologies. For the experiment, we assign the policy and wage parameters conventional values.[17]

Figure 3 compares the original allocation with full redistribution toward R. The effect on

[17]We set $\phi_\Pi = 0.5$, $\phi_Y = 0.1$, $\rho_\tau = 0$, $\phi_B = 1$, $\rho_{w,H} = \rho_{w,R} = 0.85$, $d_H = d_R = 0.25$, and $\varepsilon_{w,H} = \varepsilon_{w,R} = 1$. The remaining parameters are those in Section 5. Section 6.3 instead calibrates these parameters to the empirical multipliers.

output is large. The impact response falls from about 0.69 to 0.37 percent, and the year-2 output multiplier falls from 1.31 to 1.00. Redistribution toward the Ricardian family weakens the feedback from income to private demand relative to the original allocation: the hand-to-mouth family's additional consumption is removed, while the Ricardian family smooths the income it receives. Output therefore rises less even though the spending shock and the economy's production and labor-market structure are unchanged.

Headcount employment responds much less than output. The year-1 employment multiplier rises from 1.75 to 2.03, and the year-2 multiplier declines only slightly, from 1.69 to 1.61. The lower output response under redistribution comes from hours per worker, which rise by less because they move in proportion to the contemporaneous after-tax wage. Hiring instead depends on the value of a filled job, the expected discounted sum of the gap between the rental rate and the wage over the life of the match. Because this value is forward-looking and the spending shock is persistent, it varies little across the two allocations, and so does hiring. The model can thus generate less additional production with a similar increase in headcount employment, consistent with the empirical finding that the employment interaction is statistically insignificant while the output interaction is negative and statistically significant.[18]

6.2. A factor-demand shift and rising income concentration

The controlled exercise shows that redirecting additional wage income toward the Ricardian family lowers the output multiplier relative to the original allocation, without a comparable fall in employment. However, because it holds steady-state income concentration fixed, it does not explain why income incidence would differ across economies with different levels of inequality. We therefore study a structural change that links the two: a shift in relative factor demand toward capital and high-skill labor, which raises steady-state income concentration and, as Section 6.4 shows, changes the incidence of spending-induced income at the same time. Specifically, we lower the weight a on low-skill labor in the outer production nest Equation (4.9), holding all other parameters fixed. A lower a shifts production weight toward the capital–high-skill composite V_t, representing a shift toward capital-intensive and high-skill tasks. This interpretation is motivated by evidence that computerization reduced demand for routine tasks while increasing demand for non-routine cognitive work (Autor, Levy and Murnane, 2003), and by the role of capital–skill complementarity in

[18] This result is specific to the controlled exercise. Because the exercise leaves the production structure, and hence the composition of employment across skill groups, unchanged, hours per worker are the only margin through which output can differ from headcount. In the economies with different income concentration studied below, the composition of employment rather than hours accounts for the lower output per worker (Appendix J).

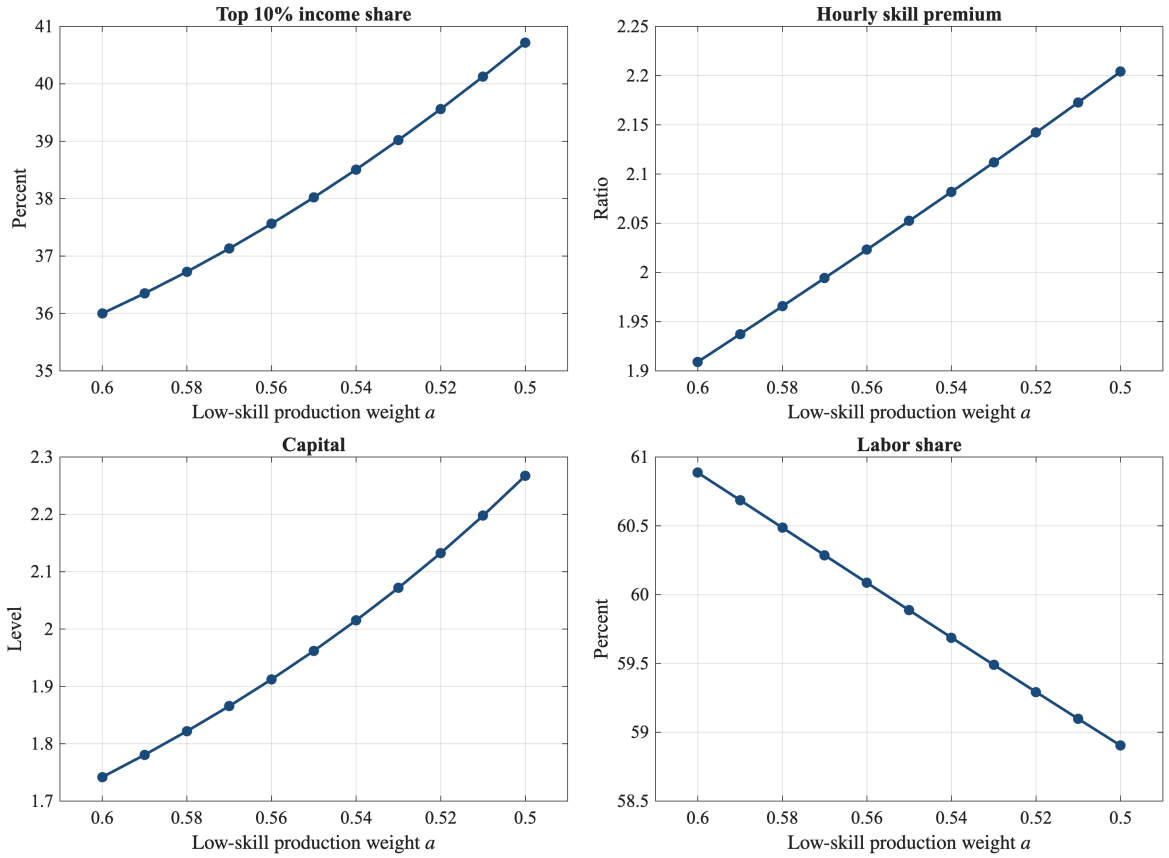


Figure 4: Steady States Across the Low-Skill Production Weight

Notes: The top 10% income share and labor share are percentages. The skill premium is the ratio of high-skill to low-skill hourly earnings. Each marker is a separately solved steady state.

explaining the skill premium (Krusell et al., 2000). The secular rise in the college wage premium and relative demand for skill documented by Autor, Katz and Kearney (2008) further motivates the direction of the comparison.

Lowering a from 0.60 to 0.50 makes production less intensive in low-skill labor and more intensive in capital and high-skill labor. Firms use more of the capital–high-skill composite relative to low-skill labor, raising its share of gross factor payments from 36 to 51 percent. The accompanying capital accumulation increases the capital stock by roughly 30 percent, while stronger demand for high-skill labor raises its wage relative to the low-skill wage. The hourly skill premium consequently rises from 1.91 to 2.20, while the aggregate labor share falls from 61 to 59 percent. Because high-skill labor and capital income accrue to the Ricardian family, these changes raise the top 10% income share from 36 to about 41 percent, an increase of 4.7 percentage points, close to the hypothetical five-percentage-point comparison in Section 2.3. Figure 4 traces these steady-state changes as a varies. Appendix G reports comparisons with other sources of income inequality.

Table 7: Dynamic Parameters Calibrated to the Empirical Multipliers

Parameter	Value	Description	Source or target
ϕ_Π, ϕ_Y	0.659, 0.118	Monetary-policy responses	Joint calibration
ρ_τ, ϕ_B	0.187, 1.076	Fiscal feedback parameters	Joint calibration
$\rho_{w,H}, \rho_{w,R}$	0.827, 0.875	Lagged-wage weights	Joint calibration
d_H, d_R	0.328, 0.003	Wage indexation weights	Joint calibration
$\varepsilon_{w,H}, \varepsilon_{w,R}$	0.984, 1.364	Wage responses to rental prices	Joint calibration

Notes: The parameters are calibrated jointly to the four moments described in the text. All other parameters are as in Section 5.

6.3. Fiscal multipliers

We next study how the response to government spending differs across these economies. We calibrate the parameters of the tax and monetary-policy rules and of wage adjustment jointly to match four moments from our main empirical specification, the pooled IV estimates: the two-year output and employment multipliers in the baseline economy and their interactions with the top 10% income share.[19] The calibrated model generates baseline multipliers of about 1.65 for output and 1.04 for employment, together with inequality interactions of -0.19 and 0.04, respectively, per percentage point of the top 10% income share. The year-2 ratios in Equation (6.1) are the model counterparts of these moments. Table 7 reports the resulting parameters. All other parameters retain the values in Section 5.

The calibrated model features substantial wage rigidity, with $\rho_{w,H} = 0.827$ and $\rho_{w,R} = 0.875$, magnitudes consistent with the importance of real-wage rigidity in search models (Hall, 2005). The calibrated monetary rule also responds to inflation by less than one-for-one ($\phi_\Pi = 0.659$), so it violates the Taylor principle. The equilibrium is nevertheless locally unique. A low ϕ_Π is also appropriate for the comparison with the data. Our empirical estimates are cross-state multipliers, for which monetary policy does not respond to state-level spending, whereas in the model the monetary rule responds to the aggregate spending shock. A weak response of the nominal interest rate to inflation, even to the extent that it violates the Taylor principle, therefore helps the model generate the substantial expansion we estimate.

With the dynamic parameters fixed, we vary a and compute the multipliers in each economy. Figure 5 plots the year-2 multipliers against the steady-state top 10% income share. The output multiplier declines steadily, from about 1.65 to 0.76. The employment multiplier follows a non-linear

[19] These aggregate moments discipline the calibration jointly but do not identify each parameter separately. We interpret the calibration as a quantitative implementation of the model rather than a structural estimate of each individual parameter.

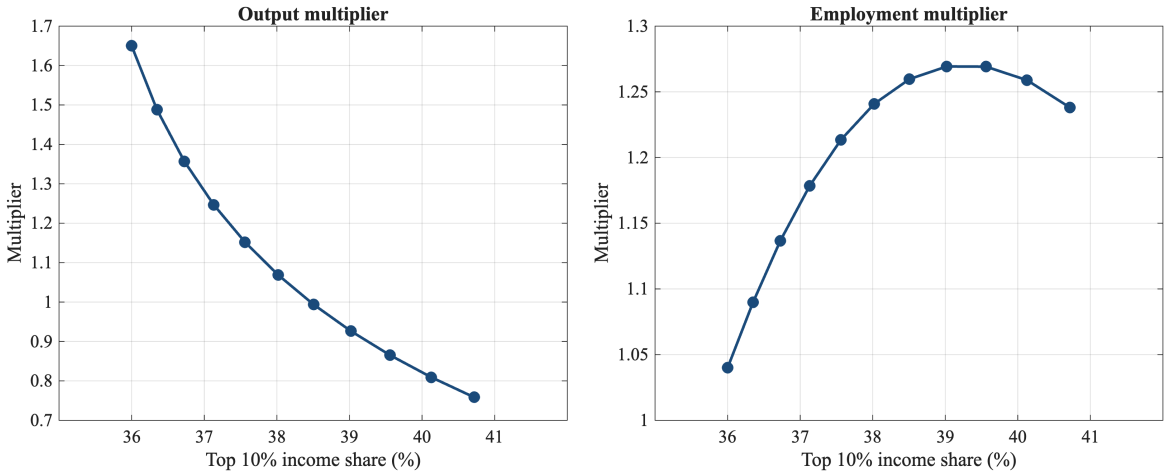


Figure 5: Year-2 Output and Employment Multipliers

Notes: Multipliers are the year-2 ratios in Equation (6.1). Each point is a separately solved economy. The shocks produce a common impact change in G/Y^{ss}.

pattern: it initially rises, peaks around $a = 0.53$, and remains above its baseline level at $a = 0.50$. The structural experiment therefore produces a lower output multiplier and a modestly higher employment multiplier, as in the pooled estimates.

To see how the responses and the gap between the two economies build up over time, Figure 6 shows the quarterly responses of the two economies and the cumulative multipliers through year five. The output response is smaller in the high-inequality economy, whereas the employment response is somewhat larger. The bottom panels use cumulative rather than annual multipliers to summarize the total response through each horizon. They differ from the year-2 ratios above, which correspond to the empirical two-year change. The gap in cumulative output multipliers between the two economies widens with the horizon, in the same direction as the local projections in Section 2.4, where the output interaction deepens over the first three years.[20]

The calibrated model thus reproduces the output-employment contrast: the same increase in government spending raises output by less in the economy with higher income concentration, while the employment response is slightly larger than in the baseline economy. The year-2 multiplier levels and inequality interactions are calibration targets. The remaining task is to explain how the model generates their contrast. We first trace who receives the income generated by additional spending, and then how the shift in factor demand changes hiring in each labor-market segment.

[20] For horizon h, these are

$$M_h^Y = \frac{\sum_{j=1}^{h}[\overline{Y}_j - Y^{ss}]}{\sum_{j=1}^{h}[\overline{G}_j - G^{ss}]}, \qquad M_h^N = \frac{\sum_{j=1}^{h}[\overline{N}_j - N^{ss}]}{\sum_{j=1}^{h}[\overline{G}_j - G^{ss}]/Y^{ss}}.$$

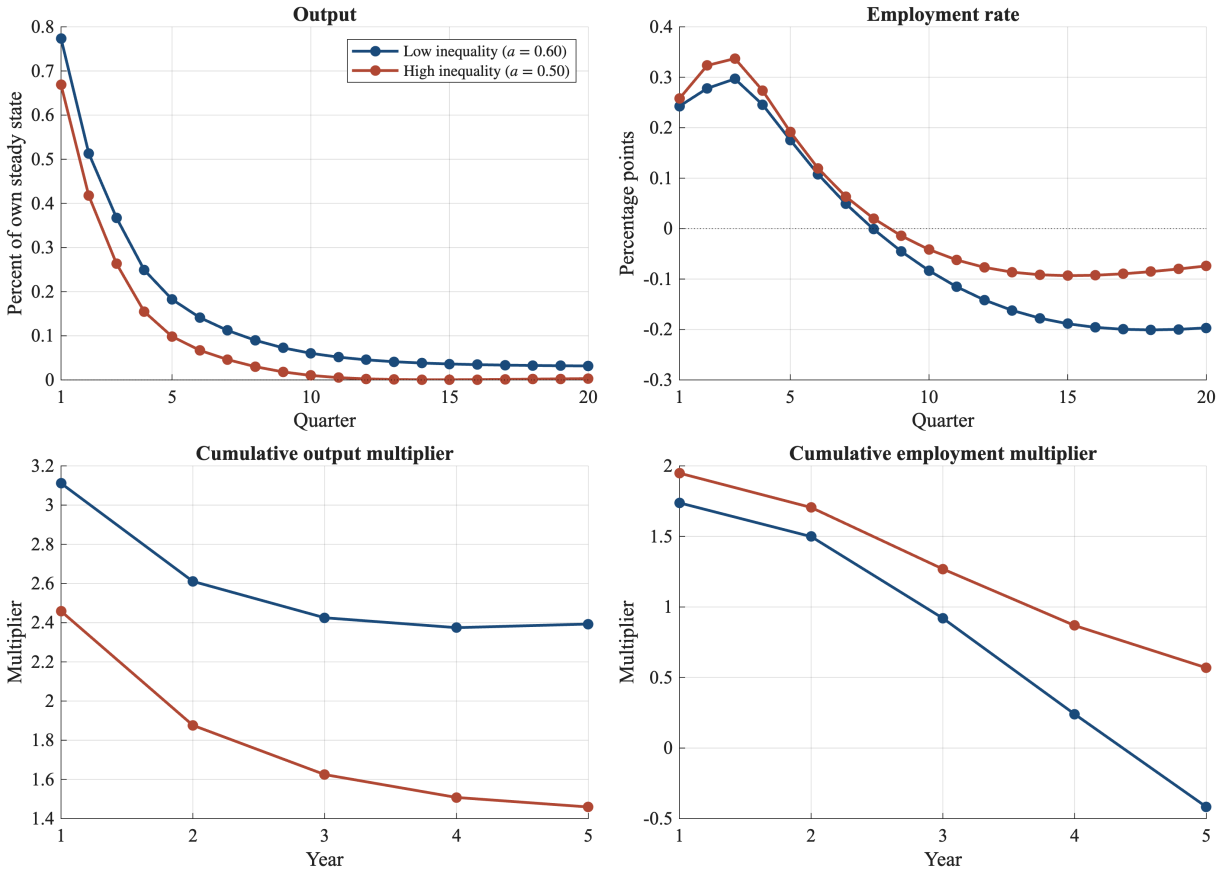


Figure 6: Quarterly Responses and Cumulative Multipliers

Notes: Low and high inequality correspond to $a = 0.60$ and $a = 0.50$. Output responses are percentage deviations from each economy's own steady state. Employment-rate responses are percentage-point deviations from their steady states. The shocks produce the same impact change in G/Y^{ss} across economies. The bottom panels show cumulative multipliers as defined in Section 6.3.

6.4. Income incidence across economies

We first examine whether the factor-demand shift redirects the income generated by a spending expansion toward the Ricardian family, as in the controlled exercise in Section 6.1. A government spending shock first raises demand for final goods, and firms respond by increasing production and their demand for inputs. In the high-inequality economy, production places greater weight on the capital–high-skill composite and less weight on low-skill labor. Producing additional output therefore draws more heavily on capital and high-skill labor relative to low-skill labor than in the baseline economy, and capital–skill complementarity further links the demand for the two inputs within the composite. The additional factor payments are consequently tilted toward high-skill workers and capital owners, whose income accrues to the Ricardian family.

Figure 7 shows this reallocation directly. In the baseline economy, the hand-to-mouth family receives about 64 percent of the impact-quarter increase in household income and the Ricardian

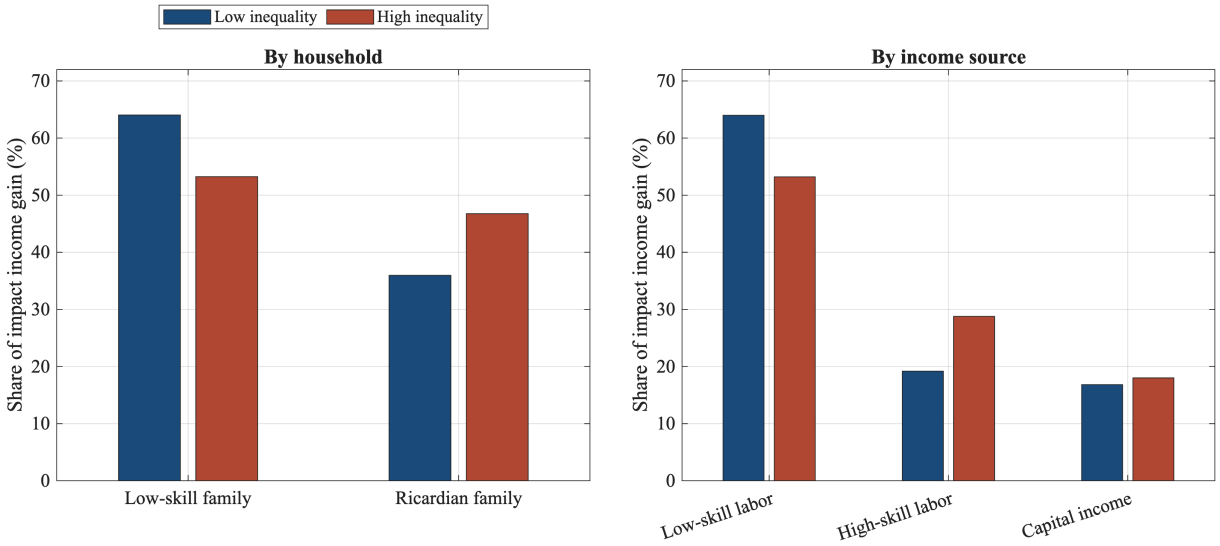


Figure 7: Shares of the Impact-Quarter Household-Income Response

Notes: Bars show deviations in the quarter of the spending shock, divided by the aggregate household-income gain in that quarter. Household shares are population-weighted. Labor-income contributions are after tax. Capital income comprises dividends, real interest on government bonds, and capital rental income net of depreciation, as in Equation (4.5). It excludes changes in the market value of existing capital claims. Source shares sum to 100 percent and may be negative.

family receives the remaining 36 percent. In the high-inequality economy, the hand-to-mouth family's share falls to about 53 percent, while the Ricardian family's share rises to about 47 percent. The decomposition by income source shows where this shift comes from. The share paid to high-skill labor rises from about 19 to 29 percent, and the capital-income share rises modestly, while the share paid to low-skill labor falls by roughly ten percentage points. The structural change therefore redirects the income gain toward the lower-MPC family, as in the controlled exercise, but without an exogenous redistribution rule.

This is the income-incidence channel isolated in Section 6.1: less of the spending-induced income reaches the hand-to-mouth family, so the feedback from income to private demand weakens. On impact, total household income per dollar of spending falls from 2.84 in the baseline economy to 2.46 in the high-inequality economy, and the income reaching HtM households falls from 1.82 to 1.31.[21] Unlike in the controlled exercise, this incidence pattern is not imposed. It is an equilibrium outcome of the same factor-demand shift that raises steady-state income concentration, so a single structural change links higher inequality to the weaker output response through the mechanism examined in Section 6.1.

[21]The bars report shares of the household-income gain, not income per dollar of government spending. The income reaching the hand-to-mouth family per dollar of spending is its share of the income gain multiplied by the total household-income gain per dollar of spending.

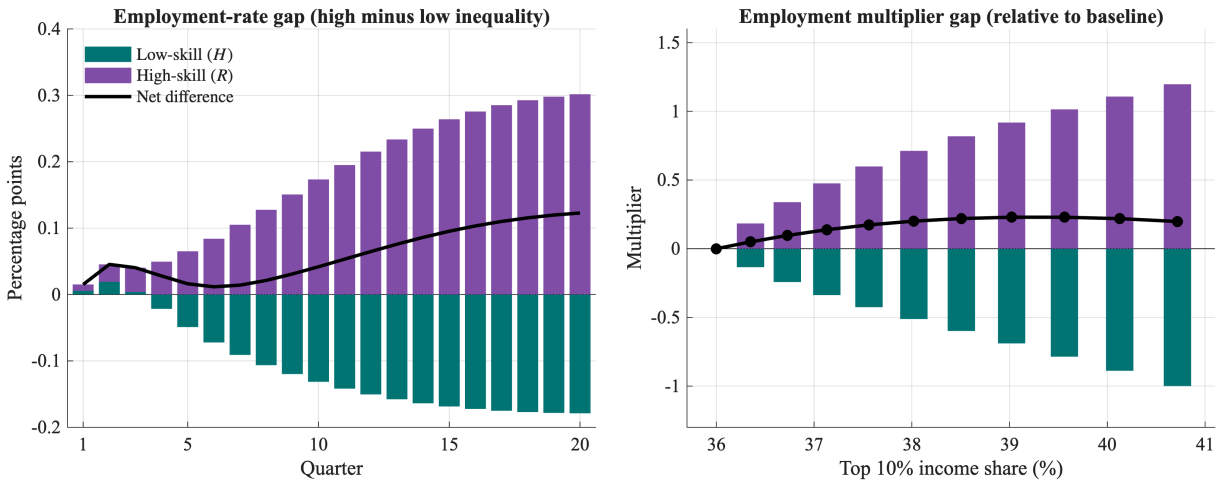


Figure 8: Employment Response Decomposition

Notes: The left panel shows the employment-rate response in the high-inequality economy ($a = 0.50$) minus that in the low-inequality economy ($a = 0.60$), in percentage points, with each response measured as a deviation from the economy's own steady state. The right panel shows the year-2 employment multiplier of each economy minus that of the low-inequality economy, with each economy plotted at its steady-state top 10% income share. In both panels, the solid line is the net difference and the bars are the contributions of low-skill (H) and high-skill (R) workers, weighted by population shares (λ and $1-\lambda$), which sum to the line. The shocks produce the same impact change in G/Y^{ss} across economies.

6.5. Employment and labor-market adjustment

Explaining the employment response also requires tracing hiring in the two labor-market segments.[22] Figure 8 decomposes the difference between the high- and low-inequality economies into the contributions of each skill type, for the employment response over time (left panel) and for the year-2 multiplier across economies with different top 10% income shares (right panel).[23] The left panel shows that the two worker types respond differently to the spending shock. The difference originates in the steady state: a lower a shifts production toward the capital–high-skill composite and away from low-skill labor. After the shock, this shift changes how labor demand responds. Relative to the baseline, demand for high-skill workers responds more, which raises aggregate employment. Demand for low-skill workers responds by less: from about the end of the first year, the employment gain of low-skill workers is smaller in the high-inequality economy than in the baseline, which lowers aggregate employment. The net difference depends on the relative size of these two contributions. The high-skill contribution is the larger one in the calibrated economy, so the employment response is larger in the high-inequality economy at every horizon, but the net

[22] Changing a alters the steady state of each segment, so we hold the conditional local elasticity of vacancy creation with respect to the rental price at its baseline value in each segment (Appendix H). The central multiplier ordering survives an alternative assumption.

[23] The employment responses themselves are in the top right panel of Figure 6.

difference is small relative to either component: at year 2 it is about 0.02 percentage points, the sum of contributions of about −0.11 from low-skill workers and +0.13 from high-skill workers.

The right panel compares the year-2 employment multiplier across economies with different steady-state top 10% shares. In economies with a higher share, the high-skill contribution is larger and the low-skill contribution smaller, but not in proportion. The net difference is therefore small and not monotonic, which is consistent with the statistically insignificant baseline employment interaction in the data. Appendix I reports the wage and unemployment responses by skill segment in each economy.

The calibration matches both the inequality interaction of the output multiplier (−0.19) and that of the employment multiplier (+0.04). Employment is therefore not what makes output rise less in the high-inequality economy. As in the empirical decomposition, the weaker output response must come from output per worker. Decomposing it into capital services, hours per worker, and the composition of employment across skill groups (Appendix J), composition accounts for the lower response: low-skill workers are about three-quarters of the employed, but their production weight is below that share and falls further when a falls, so each low-skill job adds less to output than to headcount. Hours partly offset this, whereas in the controlled exercise of Section 6.1 hours account for the lower output per worker. The difference arises because the hours term weights each group's hours response by its production weight. In the controlled exercise the weights are fixed and hours respond by less. Here the average hours response is also lower in the high-inequality economy, but production weight shifts toward high-skill workers, whose hours rise the most, and away from low-skill workers, whose hours fall below their steady-state level by year 2. This reweighting outweighs the weaker responses, so the hours term is positive. Capital services contribute little.

The experiments connect the lower output multiplier to two mechanisms within the calibrated economy. Redirecting the income generated by a spending expansion toward the Ricardian family weakens consumption feedback and lowers the output multiplier. A shift in factor demand toward capital and high-skill labor raises income concentration and redirects spending-induced income in the same direction. Segmented labor markets explain why this is not accompanied by a comparable fall in employment, and the composition of employment explains why output nevertheless rises by less.

7 Conclusion

Fiscal multipliers differ with the distribution of income. In U.S. state-level military spending data, output multipliers are smaller in states with higher top income concentration, while employment

multipliers do not differ significantly. Output per worker accounts for the difference, and personal-income gains are smaller in high-inequality states. For fiscal transmission, what matters is both the total income gain per spending dollar and the share received by households with high MPCs (Auclert et al., 2024, 2025).

The model connects these two objects through production structure. In the calibrated model, a shift in factor demand toward capital and high-skill labor raises income concentration and reduces the income that reaches hand-to-mouth (HtM) households per spending dollar. The resulting weaker consumption feedback lowers the output response. Segmented labor markets keep headcount employment from falling, and the composition of employment across skill groups means that each job adds less to output. The model thus links a lower output multiplier and a small difference in employment responses across economies to a single structural change.

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Online Appendix for

"Unequal Multipliers: Income Inequality and Government Spending"

Donggyu Lee[1] Choongryul Yang[2]

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[1]Federal Reserve Bank of New York. E-mail: donggyu.lee@ny.frb.org.
[2]Federal Reserve Bank of Atlanta. E-mail: choongryul.yang@atl.frb.org.

A Data Appendix

This appendix describes the sources and construction of the annual state-level data. The baseline panel covers all 50 states and the District of Columbia over 1966–2006, and the extended sample runs through 2014.

A.1. Outcome Variables

Real per capita output. State-level GDP (Gross State Product, or GSP) is obtained from the Bureau of Economic Analysis (BEA) regional accounts. The BEA publishes annual GDP by state for all 50 states and the District of Columbia. We deflate nominal GSP by the national Consumer Price Index for All Urban Consumers (CPI-U, BLS series CPIAUCSL) and divide by state population to obtain real per capita output. The BEA reports GSP under both SIC and NAICS classifications. We use the SIC-based series for 1966–1997 and the NAICS-based series from 1997 onward, splicing the two at the overlap year.

Employment rate. State-level nonfarm payroll employment is from the Bureau of Labor Statistics (BLS) Current Employment Statistics (CES) program. We take annual averages of the monthly establishment-survey employment counts and divide by state population from Census Bureau intercensal estimates. Employment is measured by place of work and population by place of residence.

Output per worker. Labor productivity is real GSP divided by CES payroll employment. The employment rate is CES employment divided by state population, so output per capita is their exact product. Both components are available over the 1966–2006 baseline sample.

Household employment and labor force. Household employment, unemployment, and the labor force are the BLS state estimates, available for all 51 states from 1976. Two-year differencing gives a 1978–2006 estimation window. We compute the unemployment rate from unemployment and labor-force levels because the published rate is rounded.

Manufacturing hours. We use the two-year proportional change in average weekly manufacturing hours.

State prices. The ACCRA cost-of-living series in the baseline panel is used to examine whether state consumer prices respond differently to spending across the inequality distribution.

Sectoral output and employment. Output for the ten major SIC industries is from the BEA GSP-by-industry accounts, and employment is from BEA table SA25. The output series cover 1966–1997 and the employment series cover 1969–2001. The within-industry and between-industry decomposition uses the common 1971–1997 estimation window after two-year differencing.

A.2. Spending Variable

Military spending (baseline). We use the military spending data compiled and cleaned by Nakamura and Steinsson (2014) for 1966–2006. Procurement data come from the DD-350 Individual Contract

Action Report database, which records contracts exceeding $10,000 (pre-1983) and $25,000 (post-1983). The forms identify prime contractors and the location where most of the work was performed. Our baseline spending variable is the sum of prime military contracts and military compensation. We convert military compensation from thousands of dollars to dollars before adding the two components, then deflate by the national CPI-U and divide by state population.

Military contracts (extended sample). For 1966–2014, we use the contract-only spending series from Dupor and Guerrero (2017), with 2007–2009 filled using the alternative contract series in their data. We deflate spending by the national CPI-U and divide by state population. This measure excludes military compensation.

A.3. Inequality Measures

Top income shares. The primary inequality measures are top 10%, top 5%, and top 1% income shares from the Frank-Sommeiller-Price (FSP) series (Frank, Sommeiller, Price and Saez, 2015). These are constructed from IRS *Statistics of Income* (SOI) individual tax return data, following the methodology of Piketty and Saez (2003). Pre-tax adjusted gross income includes wages, salaries, capital income (dividends, interest, capital gains), and entrepreneurial income. The FSP series covers all 50 states and DC annually from 1917 to 2015. We center each measure around its cross-state median in each year to obtain $I_{it} = \text{Top}k_{it} - \text{median}_j[\text{Top}k_{jt}]$.

Gini coefficient. We use the state-level Gini coefficients from Frank (2009) over 1966–2006. These are also based on IRS SOI data. The Gini summarizes the full income distribution, ranging from 0 (perfect equality) to 1 (maximal concentration).

Wage inequality. We use state-year weekly wage percentiles from the Current Population Survey (CPS) for 1966–2004. The wage 90/10 ratio is the 90th percentile divided by the 10th percentile of weekly wages.

Bottom-to-median income ratio. We divide the 10th percentile of income by the median within each state-year using CPS income percentiles for 1966–2004. A higher ratio means that bottom income is closer to median income. This is a measure of the income distribution, not a direct measure of liquid wealth or the HtM population share.

A.4. Other Variables

Earnings share. The BEA SA04 measure is earnings by place of work divided by earnings plus dividends, interest, and rent. The denominator measures these sources of personal income rather than GDP. Earnings include wages and salaries, supplements, and proprietors' income. This ratio is available for 1966–2006. Compensation/GDP and earnings/GDP are examined separately using the GDP-by-state accounts.

Property and proprietors' income ratio. We use BEA table SA04 for 1966–2006. The numerator is dividends, interest, rent, and proprietors' income, and the denominator is earnings plus dividends,

interest, and rent. Proprietors' income also appears within earnings, so this ratio and the earnings share are not complementary labor and capital shares. The measure concerns income received locally and does not directly measure nonlocal leakage.

Personal income components. We use BEA table SA04 for 1966–2006. Earnings by place of work comprise wages and salaries, supplements to wages and salaries, and proprietors' income. We construct wages and salaries by subtracting supplements and proprietors' income from earnings. Dividends, interest, and rent are reported as one component. Personal income equals earnings less contributions for social insurance, plus the residence adjustment, dividends, interest, and rent, and personal current transfer receipts. The series are reported in thousands of dollars. We convert them to dollars, deflate by the national CPI-U, and divide by state population.

Value-added components. For the GDP-by-state accounts, SAGDP2 reports GDP, SAGDP3 taxes on production and imports less subsidies, SAGDP4 compensation of employees, and SAGDP7 gross operating surplus. The historical SIC tables are SAGDP2S, SAGDP3S, SAGDP4S, and SAGDP7S for 1963–1997. The corresponding NAICS tables omit the S suffix and begin in 1997. We construct the component series for 1966–2006. GDP is reported in millions of dollars and the components in thousands. Both are converted to dollars before constructing the real per capita series. For GDP totals, we use NAICS levels from 1997 and SIC growth before 1997. To retain the source income shares before 1997, each state's 1997 NAICS/SIC GDP ratio is applied to all three SIC components. A separate GDP ratio is used for private industries. A small statistical discrepancy in the published accounts is retained explicitly rather than assigned to a component. This construction preserves adding-up but can leave a component-share break at 1997, so we also exclude the 1997–1998 differences and estimate the unspliced SIC and NAICS periods separately. All components are converted to real per capita dollars with the national CPI-U and state population. These GDP-component series do not replace the historical GSP outcome used in the baseline regressions.

Union density. Union density is the percentage of nonagricultural wage and salary employees who are union members. We use the UnionStats state series for 1964–2021, described by Hirsch, Macpherson and Vroman (2001), Hirsch and Macpherson (2003), and retain observations for 1966–2006.

Tax progressivity. We subtract the average state wage tax rate from the top state marginal income tax rate, both from NBER TAXSIM, and express the difference in percentage points. The series begins in 1979, giving a 1979–2006 window within the baseline sample. This variable enters as a lagged level control in Section 3.3, motivated by the tax-financing channel in Ferriere and Navarro (2025).

House price index. We use the state-level all-transactions house price index from the Federal Housing Finance Agency (FHFA). We average the quarterly observations within each year and use the log of the annual index as a control for local wealth effects. The series begins in 1975, giving a

1975–2006 window within the baseline sample.

Construction share. Construction sector GSP divided by total GSP, computed from BEA GDP-by-state data under the NAICS classification. Available from 1997 onward, yielding 510 state-year observations in the baseline sample before two-year differencing.

Education and construction shares in the productivity regressions. The college and high-school shares enter as deviations from their cross-state means in each year. The productivity regressions use these contemporaneous education shares and a state mean construction share. The latter averages the ratio of construction GSP to total GSP, using the NAICS-based ratio where available and the SIC-based ratio otherwise. These variables are interacted with spending and instrumented as described in Appendix B.2. They differ from the two-year-lagged, median-centered income shares and union density used in Section 3.3.

Population. State population comes from Census Bureau intercensal estimates. We use annual observations for all 50 states and the District of Columbia over 1966–2014 to construct per capita variables in the baseline and extended samples.

CPI. We deflate nominal monetary variables using the national Consumer Price Index for All Urban Consumers (CPI-U, all items, BLS series CPIAUCSL), with annual observations over 1966–2014. Real monetary values in the baseline panel are expressed in 2006 dollars.

A.5. Summary Statistics

Appendix Table A.1 reports summary statistics for the available observations within 1966–2006. The top 10% income share has a mean of 37.1% and a median of 36.29%, and ranges from 20.4% to 59.4%. Its within-state standard deviation is 4.59 percentage points, compared with a between-state standard deviation of 2.76 percentage points. Thus income concentration varies both across states and over time. Appendix Figure A.1 in Section A.6 displays the time-series variation by state.

Military spending averages $1,263 per capita, compared with average real per capita output of $33,587, both in 2006 dollars. The employment rate averages 41.56%. It measures CES payroll jobs divided by resident population and can exceed 100%, as in DC, because jobs are recorded by place of work. The earnings share averages 0.82 and union density 17.7%. The construction-share and tax-progressivity statistics use their shorter available periods, as reported in the table notes.

A.6. Inequality Variation

This subsection presents additional detail on the variation in income inequality across US states, which is the key source of identifying variation in the paper. Appendix Figure A.1 displays the time-series evolution of top income shares for all 51 states (including DC) from 1966 to 2006. Three patterns are notable. First, there is substantial cross-state variation in any given year: in 2006, the top 10% income share ranged from 35.8% in Alaska to 58.6% in Florida. Second, the dominant secular trend is the dramatic increase in top income concentration beginning in the early 1980s, which

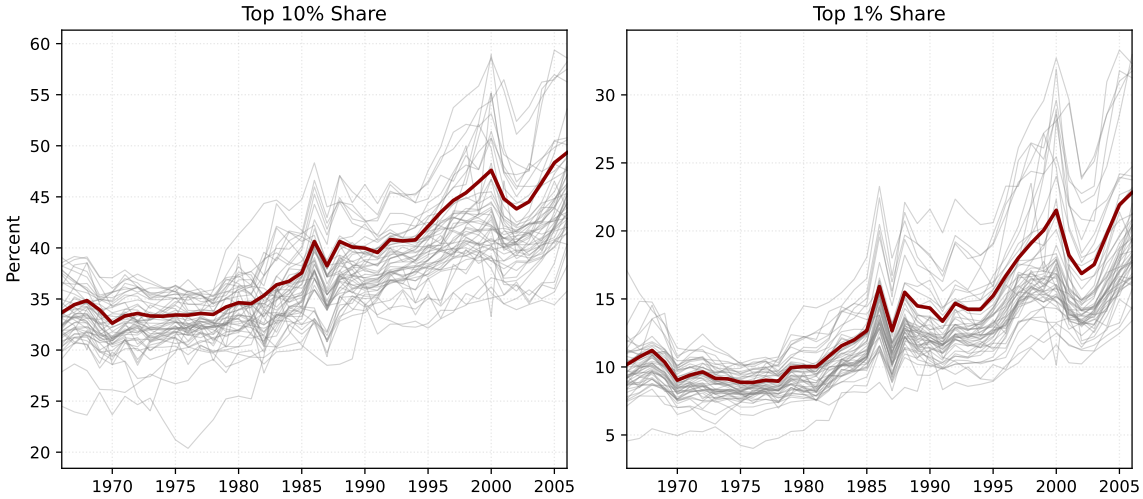
Appendix Table A.1: Summary Statistics (1966–2006)

Variable	Mean	SD	Min	Median	Max
Inequality Measures (%)					
Top 10% Income Share	37.06	5.35	20.37	36.29	59.37
Within-state SD		4.59			
Between-state SD		2.76			
Top 5% Income Share	25.78	5.26	12.65	24.64	50.19
Top 1% Income Share	12.40	4.26	4.01	11.40	33.30
Military Spending and State Variables					
Military spending (real p.c., $)	1,263	1,092	35	968	9,503
GDP by state (real p.c., $)	33,587	12,564	14,715	31,321	150,778
Employment rate (%)	41.56	10.28	22.94	40.90	117.74
Population (millions)	4.78	5.27	0.27	3.25	35.98
Mechanism Variables					
Earnings share (BEA)	0.82	0.04	0.68	0.82	0.95
Union density (%)	17.7	8.4	2.3	16.7	42.4
Construction share of GSP (%)	4.7	1.1	0.9	4.6	10.4
Tax progressivity (pp)	0.71	1.73	−8.51	0.21	14.50

Notes: Statistics use available observations within 1966–2006. The baseline panel contains 51 states, including DC, and 2,091 state-year observations before the two-year lag structure (1,989 in the baseline regressions). Construction share uses 510 observations over 1997–2006, and tax progressivity uses 1,428 over 1979–2006. The remaining level variables use 2,091 observations. Military spending is prime military contracts plus military compensation. Real monetary values are in 2006 dollars, deflated by the national CPI. Employment is CES payroll jobs divided by resident population, expressed as a percentage. This ratio can exceed 100% because jobs are measured by place of work while population is measured by residence, as in DC. SD denotes the sample standard deviation. Within-state SD is calculated after subtracting each state's mean, and between-state SD is calculated across the 51 state means. Earnings share is earnings by place of work divided by earnings plus dividends, interest, and rent. Union density is from UnionStats. Construction share is construction GSP divided by total GSP in the NAICS accounts. Tax progressivity is the top state marginal rate minus the average state wage rate, in percentage points (TAXSIM).

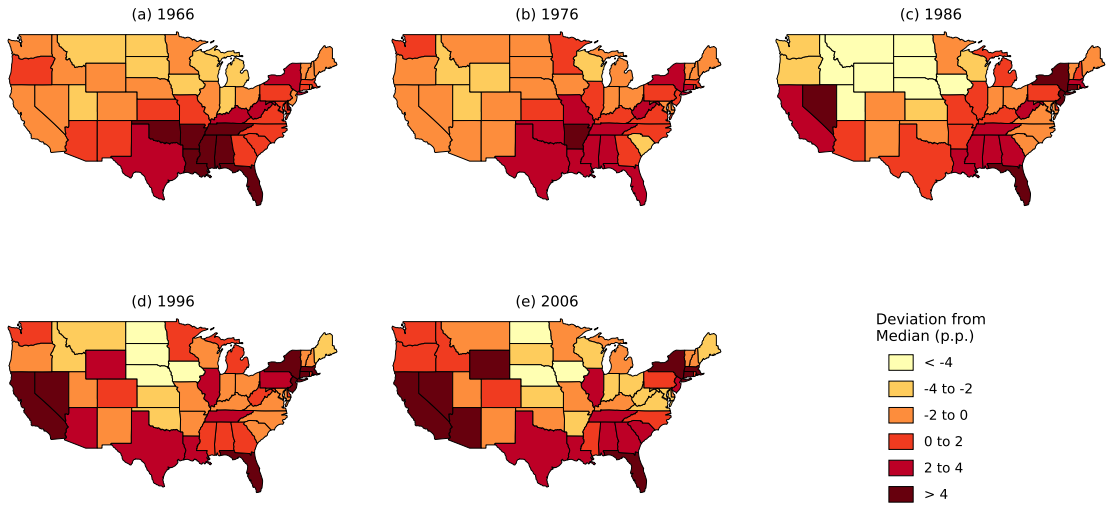
is visible in virtually every state. Third, the cross-state dispersion has widened over time: states with initially high inequality—such as Connecticut, New York, and Massachusetts—experienced the largest increases, while states with low inequality—such as West Virginia, Iowa, and Maine—saw more modest increases.

Appendix Figure A.2 displays the geographic and time-series variation of the top 10% income share. For selected years from 1966 to 2006, each state's top 10% income share is expressed as a deviation from the cross-state median for that year. The shading represents the magnitude of this deviation, with a common classification across all panels. The maps reveal that the geographic pattern of inequality has shifted over time: in the 1960s and 1970s, high-inequality states were concentrated in the South and parts of the Northeast, while by the 2000s, inequality had become particularly concentrated in states with large financial and technology sectors, such as New York, Connecticut, and Massachusetts.



Appendix Figure A.1: Historical Variation of Top Income Shares (1966–2006)

Notes: Each thin line represents one US state. The left panel shows the top 10% income share and the right panel shows the top 1% share. Bold red lines are the national averages. Source: Frank et al. (2015), based on IRS Statistics of Income data.



Appendix Figure A.2: Variation of Top 10% Income Share across States

Notes: These figures provide a visual display of top 10% income share from 1966 to 2006 for selected years. The shading of states represents the deviation (in percentage points) of the state's top 10% income share from the median value for that year, with lighter shading representing lower inequality and darker shading representing higher inequality. All panels use the same classification. Source: Frank et al. (2015), based on IRS Statistics of Income data.

B Additional Empirical Analysis

This appendix reports additional estimates and robustness checks that supplement Section 3. It includes checks on productivity responses, further income-account results, and comparisons of statistical inference.

Appendix Table B.1: Baseline and Mechanism Estimates Under Alternative Standard Errors

		State-clustered		Driscoll–Kraay	
Coefficient	Estimate	s.e.	p	s.e.	p
Panel A. Baseline estimates					
Output, no interaction	1.619	0.398	< 0.001	1.042	0.128
Output, median inequality	1.654	0.352	< 0.001	0.644	0.014
Output interaction	−0.189	0.059	0.001	0.066	0.007
Employment, no interaction	1.389	0.315	< 0.001	0.703	0.055
Employment, median inequality	1.041	0.258	< 0.001	0.449	0.026
Employment interaction	0.042	0.036	0.246	0.070	0.552
Panel B. Mechanism estimates					
Output per worker, full sample	−0.206	0.041	< 0.001	0.077	0.011
Within-industry component	−0.211	0.076	0.005	0.114	0.076
Between-industry component	0.013	0.024	0.576	0.027	0.630
Manufacturing hours	0.122	0.056	0.029	0.040	0.005
Consumer prices	−0.026	0.029	0.358	0.053	0.623
Earnings by place of work	−0.100	0.034	0.004	0.051	0.057
Wages and salaries	−0.059	0.026	0.023	0.044	0.184
Supplements	−0.013	0.005	0.018	0.008	0.127
Proprietors' income	−0.028	0.012	0.016	0.013	0.035
Dividends, interest, and rent	0.002	0.013	0.858	0.016	0.883
Compensation of employees	−0.068	0.029	0.019	0.047	0.158
Gross operating surplus	−0.100	0.043	0.020	0.031	0.002

Notes: Point estimates and estimation samples are identical across covariance estimators. Panel A reports baseline coefficients. Panel B reports spending interactions with inequality, using the samples in the corresponding tables in the main text and appendix. State-clustered p-values use the normal reference distribution. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6, and p-values use t_{T-1}, where T is the number of observed years.

B.1. Alternative Standard Errors

The panel regressions use Driscoll–Kraay standard errors to allow for heteroskedasticity, serial correlation, and dependence across states (Driscoll and Kraay, 1998). Appendix Table B.1 compares state-clustered and Driscoll–Kraay inference for the baseline estimates in Panel A and the mechanism estimates in Panel B. The regressors, instruments, samples, and coefficient estimates are identical under the two covariance estimators.

For the IV regressions, let X contain the regressors, including the fixed effects, and let Z contain the exogenous regressors and excluded instruments. We form the fitted regressors $\widehat{X} = P_Z X$ and the structural residuals $\widehat{u} = y - X\widehat{\beta}$. The covariance calculation first sums $\widehat{x}_{it}\widehat{u}_{it}$ across states within each calendar year. With $s_t = \sum_i \widehat{x}_{it}\widehat{u}_{it}$ and $A = \widehat{X}'\widehat{X}$, it is

$$\widehat{V}_{DK} = A^{-1}\widehat{M}A^{-1},$$

$$\widehat{M} = \sum_t s_t s_t' + \sum_{\ell=1}^{L}\left(1 - \frac{\ell}{L+1}\right)\sum_t \left(s_t s_{t-\ell}' + s_{t-\ell}s_t'\right).$$

The Bartlett weights decline with the calendar lag. Missing years retain their calendar spacing.

Appendix Table B.2: Local-Projection Inference Under Alternative Standard Errors and Lags

h	State p	Rule L	Rule p	$L=1$	$L=2$	$L=3$	$L=4$	$L=8$
Output inequality interaction								
0	0.001	3	0.007	0.035	0.013	0.007	0.005	0.003
1	0.023	3	0.047	0.076	0.064	0.047	0.037	0.036
2	0.051	3	0.048	0.062	0.061	0.048	0.039	0.035
3	0.073	4	0.042	0.053	0.053	0.050	0.042	0.040
4	0.276	5	0.107	0.165	0.150	0.147	0.129	0.116
5	0.889	6	0.811	0.840	0.831	0.830	0.823	0.810
6	0.967	7	0.942	0.948	0.947	0.946	0.944	0.941
7	0.813	8	0.722	0.704	0.714	0.722	0.723	0.722
8	0.790	9	0.693	0.649	0.664	0.678	0.688	0.693
Employment inequality interaction								
0	0.246	3	0.552	0.527	0.542	0.552	0.551	0.568
1	0.207	3	0.392	0.356	0.380	0.392	0.385	0.378
2	0.222	3	0.391	0.331	0.371	0.391	0.392	0.385
3	0.212	4	0.326	0.263	0.302	0.323	0.326	0.323
4	0.254	5	0.203	0.186	0.209	0.219	0.212	0.174
5	0.189	6	0.035	0.052	0.066	0.068	0.057	0.019
6	0.149	7	0.019	0.038	0.054	0.056	0.048	0.013
7	0.229	8	0.097	0.137	0.176	0.186	0.175	0.097
8	0.357	9	0.192	0.279	0.325	0.337	0.327	0.218

Notes: Entries are two-sided pointwise p-values for the inequality interaction. State-clustered inference uses the normal reference distribution. All DK columns use Bartlett weights and t_{T-1}. The rule sets $L = \max\{\lfloor 4(T/100)^{2/9} \rfloor, h+1\}$ to accommodate overlapping outcome windows. The remaining columns hold the maximum included calendar lag fixed. These are not simultaneous tests across horizons.

We apply no finite-sample covariance multiplier. For a regression with T observed years, the maximum included lag is $L = \lfloor 4(T/100)^{2/9} \rfloor$. For local projections, it is the larger of this rule and $h+1$, reflecting the overlap in changes from $t-2$ to $t+h$. We use a t_{T-1} reference distribution for tests and confidence intervals. The dynamic figures report pointwise 90% confidence bands, with inner 68% bands. The state-clustered results retain their original normal-reference inference.

Panel A shows that the baseline output inequality interaction is negative and significant under both covariance estimators, while the employment interaction is statistically insignificant under both. In Panel B, the output-per-worker and operating-surplus interactions remain negative and significant at the 5% level under Driscoll–Kraay inference. The earnings and within-industry interactions are significant at the 10% level, with p-values of 0.057 and 0.076, respectively. The compensation interaction has a p-value of 0.158. Compared with state-clustered standard errors, Driscoll–Kraay standard errors are larger for some estimates and smaller for others.

Appendix Table B.2 compares local-projection inference under state clustering, the horizon-dependent Driscoll–Kraay lag rule, and fixed maximum lags of one, two, three, four, and eight years. The output inequality interaction is significant at the 10% level through $h = 3$ under state clustering and each Driscoll–Kraay lag choice. Under the horizon-dependent rule, it is significant at the 5%

Appendix Table B.3: Additional Checks on Productivity Responses

Specification	Dependent variable	$\hat{\beta}_d$	s.e.	N
Panel A. State characteristics, dependent variable $\Delta\log(Q/E)$				
Inequality only		-0.206^{**}	(0.077)	1,989
+ college share		-0.173^{***}	(0.057)	1,989
+ high-school share		-0.213^{***}	(0.077)	1,989
+ construction share		-0.237^{***}	(0.054)	1,989
+ all three		-0.217^{***}	(0.043)	1,989
Panel B. Hours and prices				
Intensive margin	Δ hours, manufacturing	0.122^{***}	(0.040)	1,438
Deflation	Δ state CPI (ACCRA)	-0.026	(0.053)	1,763

Notes: Each row is a separate IV regression using the baseline design. $\hat{\beta}_d$ is the interaction of spending with the lagged top-10% share deviation. Panel A uses the full baseline sample, 1968–2006. Each additional state characteristic enters interacted with spending and is separately instrumented. Time-invariant characteristics are absorbed by the state fixed effects and enter through the interaction only. Appendix A.4 describes the education and construction shares, and Appendix A.1 describes the hours and price series. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years. $^{***}p<0.01$, $^{**}p<0.05$, $^{*}p<0.1$.

level through $h=3$.

The employment interaction is statistically insignificant at every horizon under state clustering. Under Driscoll–Kraay inference, the 90% bands exclude zero at $h=5$ and $h=6$ under each fixed lag choice. Significance at the 5% level depends on the lag choice: the p-values are 0.035 and 0.019 under the horizon-dependent rule, compared with 0.068 and 0.056 with three lags. The $h=7$ interaction has a p-value of 0.097 under the horizon-dependent rule and 0.186 with three lags. These significance comparisons refer to individual horizons.

The baseline contains 39 years, while several sectoral and skill-specific samples contain eight. The small number of years can limit the accuracy of the Driscoll–Kraay approximation, even with a t reference distribution.

B.2. Differences in Productivity Responses

This subsection examines additional explanations for the smaller output-per-worker response in high-inequality states. Appendix Table B.3 reports spending interactions with education and construction shares, along with checks using manufacturing hours and state consumer prices. Appendix Table B.4 separates changes within industries from changes in employment shares across industries.

Other state characteristics. Inequality correlates with education and industrial structure. Adding spending interactions with the state college share, high-school share, and construction share—each separately instrumented—leaves the inequality interaction negative, ranging from -0.173 to -0.237 and significant at the one percent level in every augmented specification. All specifications use the same 1,989 observations. These comparisons show that the inequality interaction survives the

Appendix Table B.4: Within-Industry and Between-Industry Productivity Responses

Component	Dependent variable	$\hat{\beta}_d$	s.e.	N
Total	Δ productivity	−0.197	(0.117)	1,372
Within-sector		−0.211*	(0.114)	1,372
Between-sector		0.013	(0.027)	1,372

Notes: IV estimates using the baseline design, 1971–1997. $\hat{\beta}_d$ is the interaction of spending with the lagged top-10% share deviation. The rows decompose the two-year change in state labor productivity across the ten major SIC sectors using midpoint weights. The two components add up exactly to the total before rounding (decomposition residual 7×10^{-16}). Sectoral GSP and employment use consistent SIC definitions. Appendix A.1 describes the series. State and year fixed effects. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

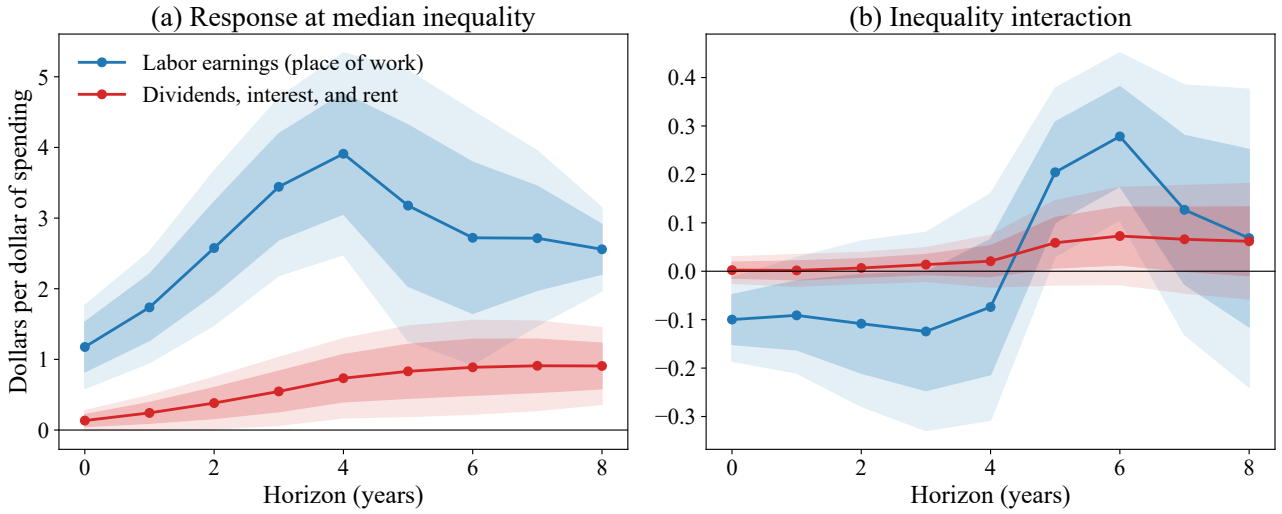
additional spending interactions, while leaving its structural source unresolved.

Within-industry and between-industry responses. We next examine whether changes in employment shares across industries account for the smaller output-per-worker response in high-inequality states. We use an exact shift-share decomposition on the ten major SIC sectors, which sum to total GSP. Writing productivity as $P_{it} = \sum_j s_{ijt} p_{ijt}$, with employment shares s_{ij} and sector productivity p_{ij}, midpoint weights divide the two-year change into within-sector and between-sector terms. The output and employment accounts restrict the estimation window to 1971–1997, compared with 1968–2006 for the main productivity decomposition.

Appendix Table B.4 reports the decomposition. The within-industry interaction is −0.211 (s.e. 0.114), significant at the 10% level, and the between-industry interaction is 0.013 (s.e. 0.027). Their sum before rounding gives the total productivity interaction of −0.197 (s.e. 0.117), which is statistically insignificant. The point estimates place the smaller productivity response in high-inequality states predominantly within broad industries. The shorter sample and uncertainty surrounding the estimates limit the strength of this conclusion. The between-industry term measures changes in employment shares across broad industries. Differences in states' initial industry composition also enter the within-industry term through its weights. Differences in hours, capital services, and worker composition within industries remain unresolved.

Intensive margin. We measure manufacturing hours by average weekly hours. Because the main analysis measures output per worker, differences in hours responses could account for differences in productivity responses. The manufacturing-hours interaction is +0.122 (s.e. 0.040), so hours respond more in high-inequality states. This response works against an explanation based on smaller hours gains in manufacturing. Hours responses outside manufacturing remain unresolved because comparable state data are unavailable.

Deflation. Using the ACCRA state cost-of-living series, the price interaction is −0.026 (s.e. 0.053) and statistically insignificant. This series measures consumer prices. Differences in output prices remain a possible explanation because the consumer-price series provides an incomplete measure



Appendix Figure B.1: Local Projection Impulse Responses: Labor Earnings and Capital Income

Notes: Local projection IV estimates at horizons $h = 0, \ldots, 8$ for earnings by place of work and for dividends, interest, and rent, both from the BEA state personal income accounts. The outcome at horizon h is $(X_{i,t+h} - X_{i,t-2})/Y_{i,t-2}$, so coefficients are dollars per dollar of spending. Panel (a) reports the response at median inequality, $\hat{\beta}_m^h$. Panel (b) reports the inequality interaction, $\hat{\beta}_d^h$. Shaded areas represent 68% (inner, darker) and 90% (outer, lighter) pointwise confidence bands based on Driscoll–Kraay standard errors and t_{T-1} critical values. The maximum lag is $\max\{\lfloor 4(T/100)^{2/9} \rfloor, h+1\}$, where T is the number of years in the estimation sample.

of the prices relevant for state production.

B.3. Dynamics of the Incidence Response

Appendix Figure B.1 reports local projection IV estimates, as in Equation (2.3), for earnings by place of work and dividends, interest, and rent. The outcome at horizon h is the change in the real per capita component from $t-2$ to $t+h$ divided by real per capita output at $t-2$. The earnings response at median inequality builds to about four dollars per dollar of spending at $h = 4$. Its inequality interaction is negative through $h = 4$, although the 90% confidence band excludes zero only at $h = 0$ within this interval. The interaction becomes positive, with the 90% band excluding zero at $h = 5$ and $h = 6$. The response of dividends, interest, and rent is small at every horizon and its interaction is indistinguishable from zero throughout. These are pointwise comparisons across horizons.

B.4. Income shares in the horse races

Appendix Table B.5 compares the earnings share used in Table 4 with compensation and earnings divided by GDP. All three specifications use the same 1,989 observations and reestimate the baseline on the augmented regression's sample. Compensation/GDP does not attenuate the inequality coefficient. Earnings/GDP attenuates it by about 8%. The 39% attenuation for the earnings share therefore does not establish a general labor-share explanation. Each measure is a lagged income share, not the income response per spending dollar.

Appendix Table B.5: Horse Races with Alternative Income Shares

Additional spending interaction	$\hat{\beta}_d$	s.e.	Attenuation	Interaction with share
Earnings / (earnings + property income)	−0.115	(0.086)	39.3%	6.37 (10.68)
Compensation / GDP	−0.204	(0.082)	-7.9%	12.87 (8.82)
Earnings / GDP	−0.174	(0.068)	7.9%	−4.40 (8.38)

Notes: The baseline interaction on the common sample is −0.189 (s.e. 0.066). Each share is lagged two years and centered on its year median, and enters both directly and interacted with spending. Spending interactions are instrumented with the corresponding state-specific national spending interactions. Earnings include proprietors' income. Property income denotes dividends, interest, and rent. GDP is from the reconstructed BEA component accounts. Attenuation is calculated from unrounded coefficients and can be negative. State and year fixed effects. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years.

B.5. Additional Control Variables

We verify that the inequality interaction is not absorbed by several additional control variables, each entered as a lagged level in the baseline inequality-interaction specification. Appendix Table B.6 reports the results.

Appendix Table B.6: Robustness to Additional Controls

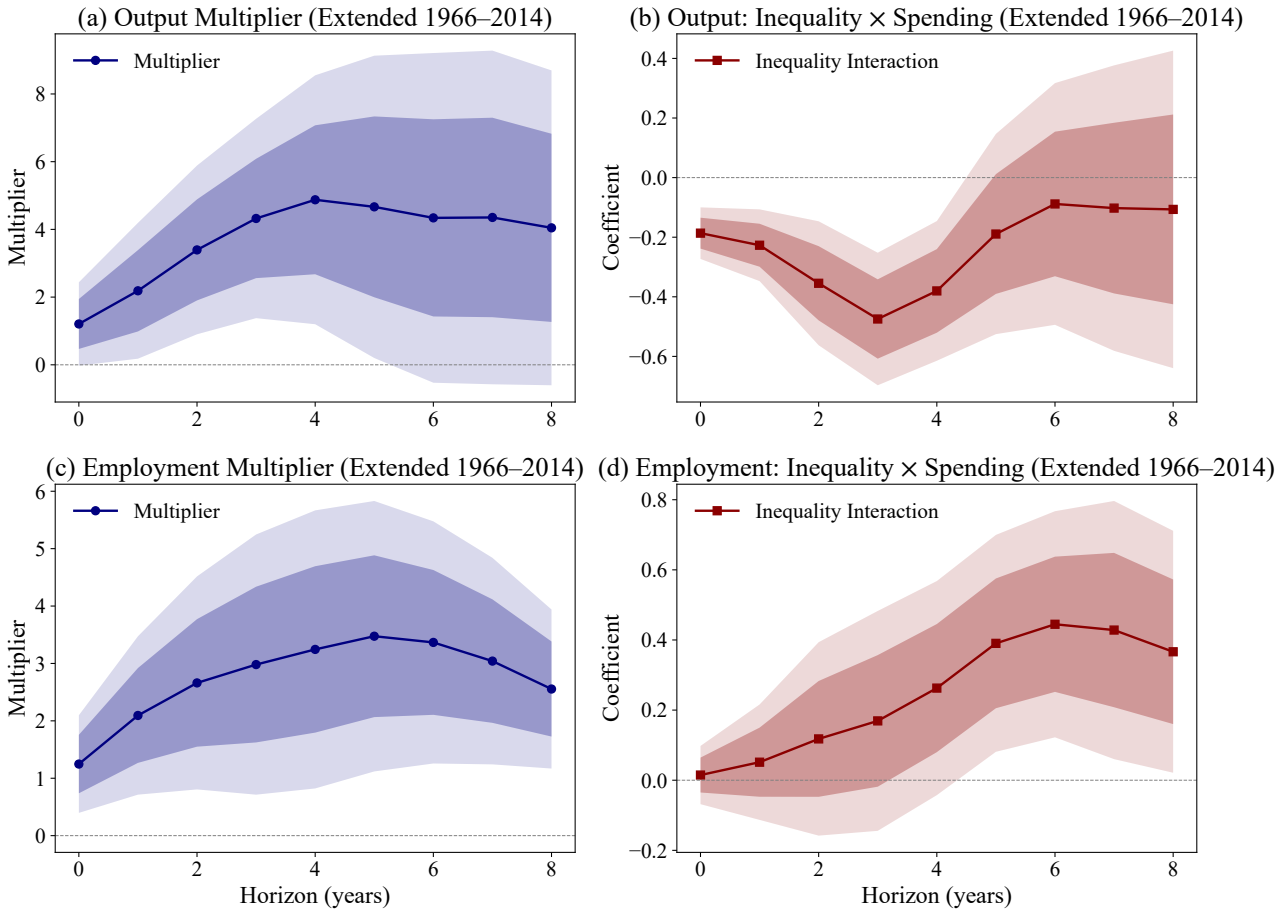
Control added	$\hat{\beta}_m$	(s.e.)	$\hat{\beta}_d$	(s.e.)	N
None (baseline)	1.65	(0.64)	−0.189	(0.066)	1,989
Earnings share	1.64	(0.66)	−0.190	(0.067)	1,989
Property income ratio	1.55	(0.61)	−0.181	(0.070)	1,989
Union density	1.66	(0.66)	−0.189	(0.066)	1,989
House price level	1.10	(0.61)	−0.187	(0.094)	1,530

Notes: Each row adds the indicated lagged control variable to the baseline inequality-interaction specification (column 2 of Table 1). Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years.

The inequality interaction is virtually unchanged when controlling for the earnings share (−0.190), the property and proprietors' income ratio (−0.181), union density (−0.189), or house price levels (−0.187). These variables enter as lagged level controls. The spending-interaction specifications in Section 3.3 additionally allow the spending response to vary with the state characteristic. The estimates here show that the inequality interaction changes little with the additional controls and leave its structural source unresolved.

B.6. Extended Sample Local Projections

This appendix presents local projection impulse responses for the extended sample (1966–2014) using contract-only spending from Dupor and Guerrero (2017). The extended sample adds eight years of data (2007–2014), which includes the Great Recession, the subsequent recovery, and a period of continued fiscal consolidation.



Appendix Figure B.2: LP-IRF Output and Employment: Extended Sample (1966–2014)

Notes: Local projection IV estimates using contract-only spending from Dupor and Guerrero (2017) over 1966–2014. Top row: output (panels (a) and (b)). Bottom row: employment (panels (c) and (d)). Left column: spending coefficient from the specification without the inequality interaction. Right column: inequality interaction $\hat{\beta}_d^h$ from Equation (2.3). Shaded areas represent 68% (inner, darker) and 90% (outer, lighter) pointwise confidence bands based on Driscoll–Kraay standard errors and t_{T-1} critical values. The maximum lag is $\max\{\lfloor 4(T/100)^{2/9} \rfloor, h+1\}$, where T is the number of years in the estimation sample.

Appendix Figure B.2 presents the LP-IRFs for both output and employment. The output inequality interaction is negative and significant at the 5% level from $h = 0$ through $h = 4$, and is imprecisely estimated thereafter. The employment interaction is not statistically significant at short horizons. It becomes positive, with the 90% confidence band excluding zero from $h = 5$ through $h = 8$, and the 95% band excluding zero at $h = 5$ and $h = 6$. As in the baseline sample, the employment comparison therefore differs between short and longer horizons. The maximum lag in the Driscoll–Kraay covariance increases with the horizon to accommodate overlapping outcome changes. Appendix B.1 reports the sensitivity of the baseline employment interactions to this choice.

Appendix Table B.7: Spending Definitions on a Common Sample

Spending measure	Outcome	$\hat{\beta}_m$	s.e.	$\hat{\beta}_d$	s.e.
Contracts and compensation	Output	1.654	(0.644)	−0.189	(0.066)
Contracts only	Output	1.588	(0.551)	−0.259	(0.094)
Contracts and compensation	Employment	1.041	(0.449)	0.042	(0.070)
Contracts only	Employment	1.039	(0.367)	0.010	(0.063)

Notes: All specifications use 1,989 observations over 1968–2006 and include state and year fixed effects. Output is a proportional change in real per capita GSP. Employment is a proportional change in CES payroll employment divided by population. Spending is divided by lagged real per capita output in both cases. Driscoll–Kraay standard errors use Bartlett weights and the lag rule in footnote 6. Significance uses t_{T-1}, where T is the number of observed years.

B.7. Military contracts and compensation

The baseline spending measure includes military compensation as well as contracts. Appendix Table B.7 compares it with contracts alone on the same years and states. The national spending series and both sets of instruments are reconstructed for each measure. The negative output interaction remains with contracts only. Neither spending definition gives a significant employment interaction. This comparison holds the period fixed, unlike the extended-sample exercise.

C Aggregate Evidence Using Ramey News Shocks

To complement the state-level analysis, we provide aggregate time-series evidence using the Ramey (2011) defense news shock series. This appendix tests whether the inequality–multiplier relationship documented at the state level also appears in aggregate US data using a fundamentally different identification strategy.

C.1. Data and Methodology

We use historical US data on GDP, the GDP deflator, government spending, federal receipts, and inequality. We employ annual data since the aggregate US inequality series is available only at annual frequency. The identified government spending shock is the defense news variable constructed by Ramey (2011), which is based on narrative records and measures the present value of new military spending at the time of announcement. For inequality, we use two measures: the top 10% *income* share and the top 10% *wealth* share, the latter from Saez and Zucman (2016). The sample period is 1929–2013.

We transform real GDP and real government spending using the Gordon and Krenn (2010) transformation, as in Ramey (2016). We divide both series by trend real GDP, obtained by fitting a quadratic time trend to log real GDP and exponentiating the fitted values. This expresses output and spending in the same units, so that the multiplier can be estimated directly. The tax control is federal receipts divided by GDP. We normalize the defense news shock by lagged nominal trend GDP.

To calculate impulse responses and cumulative multipliers, we use Jordà's (2005) local projection method. We estimate a set of regressions for each horizon h:

$$x_{t+h} = I_{t-1}\left[\alpha_{H,h} + \gamma_{H,h}\,\text{shock}_t\right] + (1 - I_{t-1})\left[\alpha_{L,h} + \gamma_{L,h}\,\text{shock}_t\right] + I_{t-2}\psi'_{H,h} z_{t-1} + (1 - I_{t-2})\psi'_{L,h} z_{t-1} + \tau_{1,h} t + \tau_{2,h} t^2 + \varepsilon_{t+h}.$$

where x is the variable of interest, z_t contains transformed real GDP, transformed real government spending, the tax rate, and the news shock, and $\psi_{H,h}$ and $\psi_{L,h}$ are coefficient vectors. We include one lag of these controls. The intercept and contemporaneous shock coefficient depend on I_{t-1}, while the lagged controls are interacted with I_{t-2}. The linear and quadratic trend coefficients are common across high- and low-inequality periods. Since both inequality series have trends over the sample period, we detrend their logarithms using a quadratic trend and define $I_t = 1$ if the detrended measure is above zero. As an alternative, we remove the trend with the Hodrick–Prescott filter using the annual smoothing parameter of 6.25.

Cumulative multipliers are obtained via the instrumental variable approach of Ramey and Zubairy (2014):

$$\sum_{j=0}^{h} y_{t+j} = I_{t-1}\left[\delta_{H,h} + \beta_{H,h}\sum_{j=0}^{h} g_{t+j}\right] + (1 - I_{t-1})\left[\delta_{L,h} + \beta_{L,h}\sum_{j=0}^{h} g_{t+j}\right] + I_{t-2}\phi'_{H,h} z_{t-1} + (1 - I_{t-2})\phi'_{L,h} z_{t-1} + \kappa_h t + \epsilon_{t+h}.$$

using $I_{t-1} \times \text{shock}_t$ and $(1 - I_{t-1}) \times \text{shock}_t$ as instruments for the respective interactions of cumulative government spending with the two inequality indicators. The estimated coefficients $\beta_{H,h}$ and $\beta_{L,h}$ are the cumulative multipliers at horizon h in high- and low-inequality periods, respectively. The cumulative regressions with separate high- and low-inequality multipliers include a common linear trend. The linear model without inequality interactions includes both linear and quadratic trends and one lag of each control. Standard errors are Newey–West with a Bartlett kernel and bandwidth $h+1$, following Ramey and Zubairy (2018), and the test of equal multipliers uses the full covariance matrix of $(\beta_{H,h}, \beta_{L,h})$.

C.2. Results

Appendix Table C.1 presents the cumulative multipliers for 2-year and 4-year horizons. The linear model (without inequality state-dependence) yields cumulative multipliers of 1.12 and 1.02 at the 2-year and 4-year horizons, respectively, both significantly different from zero.

When we split by income inequality detrended using a quadratic trend, the 2-year cumulative multiplier is 3.35 in low-inequality periods versus 0.90 in high-inequality periods, and the 4-year multiplier is 2.67 versus 0.84. The difference is statistically significant at the 5% level at the 2-year horizon ($p = 0.02$) and at the 1% level at the 4-year horizon ($p < 0.01$).

When we use the top 10% wealth share instead of the income share, the pattern is similar but weaker: the low-inequality multiplier exceeds the high-inequality multiplier at both horizons

Appendix Table C.1: Cumulative Multipliers Using Aggregate Data

		1. Top 10% *Income* Share			2. Top 10% *Wealth* Share		
	Linear Model	High (β_H)	Low (β_L)	p (H_0: =)	High (β_H)	Low (β_L)	p (H_0: =)
Quadratic trend							
2-year integral	1.12***	0.90***	3.35***	0.02	1.01***	1.23***	0.47
	(0.12)	(0.26)	(1.10)		(0.17)	(0.20)	
4-year integral	1.02***	0.84***	2.67***	0.00	0.89***	1.19***	0.09
	(0.08)	(0.06)	(0.39)		(0.07)	(0.14)	
HP filter							
2-year integral	1.12***	1.03***	1.53***	0.11	0.48	1.18***	0.71
	(0.12)	(0.13)	(0.32)		(1.96)	(0.19)	
4-year integral	1.02***	0.98***	1.80***	0.03	0.88***	1.17***	0.23
	(0.08)	(0.09)	(0.42)		(0.16)	(0.13)	

Notes: Each cell reports a cumulative multiplier at the indicated horizon from the one-step IV local projection, using the Ramey (2011) news shock as the instrument. "Linear model" indicates multipliers estimated without the inequality interaction. Inequality is high ($I_{t-1} = 1$) if the log top 10% share exceeds its trend, removed with a quadratic trend (upper panel) or the Hodrick–Prescott filter with smoothing parameter 6.25 (lower panel). Standard errors in parentheses are Newey–West with a Bartlett kernel and bandwidth $h+1$. The p-values test the null hypothesis that the high- and low-inequality multipliers are equal, using the full covariance matrix. Annual data, 1929–2013 ($N = 81$ at the 2-year and 79 at the 4-year horizon). $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

(1.23 versus 1.01 and 1.19 versus 0.89), but the difference is not statistically significant at the 2-year horizon and only marginally significant at the 4-year horizon ($p = 0.09$ under the quadratic trend). Detrending with the HP filter instead of a quadratic trend yields qualitatively similar results for income inequality (1.53 versus 1.03 at two years and 1.80 versus 0.98 at four years), with the difference significant at the 4-year horizon ($p = 0.03$). The wealth-share differences are insignificant under the HP filter.

D The Effective-Exposure Decomposition

This appendix presents the household-income decomposition used to interpret the model. In economy i, a fraction $\widetilde{\lambda}_i$ of households is HtM and the remaining fraction is Ricardian. Let $\Delta y_{H,i}^{disp}(h)$ and $\Delta y_{R,i}^{disp}(h)$ be population-weighted disposable-income responses at horizon h, so that the total household disposable-income response is $\Delta y_i^{disp}(h) = \Delta y_{H,i}^{disp}(h) + \Delta y_{R,i}^{disp}(h)$. Income reaching HtM households per spending dollar satisfies

$$\phi_i(h) \equiv \frac{\Delta y_{H,i}^{disp}(h)}{\Delta G_i(h)} = s_{H,i}(h)\frac{\Delta y_i^{disp}(h)}{\Delta G_i(h)}, \qquad s_{H,i}(h) \equiv \frac{\Delta y_{H,i}^{disp}(h)}{\Delta y_i^{disp}(h)}.$$

All income and spending responses use the same horizon and aggregation. The factorization requires a nonzero total income response. The income share $s_{H,i}(h)$ generally differs from the

population share $\widetilde{\lambda}_i$, since disposable-income responses per household can differ across types. Income generated per spending dollar can exceed one, and response shares can lie outside $[0,1]$ if some components decline.

The analogous labor-income calculation uses the HtM share of incremental payroll, with taxes and transfers accounted for separately. Profits or imported inputs are not additional deductions from payroll already measured as local labor income. A smaller HtM income response can reflect a smaller total income gain, a smaller HtM share, or both.

The effect on consumption depends on recipients' MPCs and the timing of income. The output multiplier additionally depends on investment, prices, financing, and other equilibrium responses. This identity does not impose a lower bound of one on the multiplier or identify a universal mapping from ϕ_i to the regression coefficients. The quantitative model specifies those additional responses. The empirical accounts identify income sources, not recipients' MPCs. The horse races do not identify $\widetilde{\lambda}_i$ or establish that it is constant across states.

Each model outcome must be measured using the empirical annual change and spending denominator. Baseline coefficients describe output and proportional employment-rate responses at median lagged inequality and their interactions with a one-percentage-point deviation in the top 10% share. Income-incidence coefficients use each income component's change divided by lagged output. These distinctions are necessary before comparing model response shares with empirical dollars per spending dollar.

E Model details

This appendix collects the model equations that are not needed to follow the income-incidence argument in Section 4 and describes the solution method.

E.1. Ricardian household optimality

Depreciation depends on utilization according to

$$\delta(\nu_t) = \delta_0 + \delta_1 \nu_t^{\delta_2}. \tag{E.1}$$

The Ricardian family's stochastic discount factor is

$$\Lambda_{t,t+1} = \beta \left(\frac{X_{R,t+1}}{X_{R,t}} \right)^{-\sigma}, \tag{E.2}$$

and optimal bond and capital holdings satisfy

$$1 = \mathbb{E}_t \left[\Lambda_{t,t+1} \frac{R_t}{\Pi_{t+1}} \right], \tag{E.3}$$

$$Q_t = \mathbb{E}_t \left\{ \Lambda_{t,t+1} \left[Q_{t+1} + r^k_{t+1} \nu_{t+1} - \delta(\nu_{t+1}) \right] \right\}. \tag{E.4}$$

The corresponding realized returns are

$$R^B_{t+1} = \frac{R_t}{\Pi_{t+1}}, \qquad R^K_{t+1} = \frac{Q_{t+1} + r^k_{t+1} v_{t+1} - \delta(v_{t+1})}{Q_t}. \tag{E.5}$$

Equation (E.3) and (E.4) imply the risk-adjusted no-arbitrage condition

$$\mathbb{E}_t \left[\Lambda_{t,t+1} \left(R^K_{t+1} - R^B_{t+1} \right) \right] = 0, \tag{E.6}$$

which ties the expected return on installed capital to the return on bonds. The price Q_t adjusts so that the Ricardian family is willing to hold the outstanding capital stock.

The owner of installed capital also chooses utilization. Its first-order condition equates the marginal rental benefit to the marginal depreciation cost,

$$r^k_t = \delta'(v_t) = \delta_1 \delta_2 v_t^{\delta_2 - 1}. \tag{E.7}$$

E.2. Price setting

Intermediate-goods producers are monopolistically competitive and face Rotemberg price-adjustment costs (Rotemberg, 1982),

$$\mathscr{A}^p_t = \frac{\phi_p}{2} \left(\frac{\Pi_t}{\bar{\Pi}} - 1 \right)^2 Y_t. \tag{E.8}$$

Let η denote the elasticity of substitution across differentiated goods. In a symmetric equilibrium, optimal price setting implies the Phillips curve

$$\begin{aligned} \left(\frac{\Pi_t}{\bar{\Pi}} - 1 \right) \frac{\Pi_t}{\bar{\Pi}} = \mathbb{E}_t \left[\Lambda_{t,t+1} \frac{Y_{t+1}}{Y_t} \left(\frac{\Pi_{t+1}}{\bar{\Pi}} - 1 \right) \frac{\Pi_{t+1}}{\bar{\Pi}} \right] \\ + \frac{\eta}{\phi_p} \left(\mathrm{MC}_t - \frac{\eta - 1}{\eta} \right). \end{aligned} \tag{E.9}$$

E.3. Capital production, profits, and dividends

A competitive capital-producing sector transforms final goods into installed capital. Let $x_{k,t} = K_{t+1}/K_t$ and let the capital-growth adjustment cost be

$$\mathscr{A}^K_t = \frac{\phi_K}{2} \left[\log(x_{k,t}) \right]^2 K_{t+1}. \tag{E.10}$$

The capital producer's flow profit is

$$\mathscr{P}^K_t = Q_t (K_{t+1} - K_t) - (K_{t+1} - K_t) - \mathscr{A}^K_t, \tag{E.11}$$

and its optimality condition determines the supply price of installed capital,

$$\begin{aligned} Q_t = 1 + \phi_K \log(x_{k,t}) + \frac{\phi_K}{2} \left[\log(x_{k,t}) \right]^2 \\ - \mathbb{E}_t \left\{ \Lambda_{t,t+1} \left[1 + \phi_K \log(x_{k,t+1}) x_{k,t+1} - Q_{t+1} \right] \right\}. \end{aligned} \tag{E.12}$$

Together, the capital-supply condition (E.12) and the household capital-demand condition (E.4) determine Q_t and capital accumulation. Physical investment is

$$I_t = K_{t+1} - [1 - \delta(\nu_t)]K_t. \tag{E.13}$$

Aggregate pre-tax profit is the sum of profits earned in goods production, labor intermediation, and capital production,

$$\mathscr{P}_t = \mathscr{P}_t^Y + \mathscr{P}_t^L + \mathscr{P}_t^K, \tag{E.14}$$

$$\mathscr{P}_t^Y = Y_t(1 - \mathrm{MC}_t) - \mathscr{A}_t^p - \Phi, \tag{E.15}$$

$$\begin{aligned}\mathscr{P}_t^L = {} & (r_{H,t} - w_{H,t})L_{H,t} + (r_{R,t} - w_{R,t})L_{R,t} \\ & - \lambda\iota_H V_{H,t} - (1-\lambda)\iota_R V_{R,t},\end{aligned} \tag{E.16}$$

where Φ is a fixed operating cost. Aggregate dividends distributed to the Ricardian family are

$$D_t = (1 - \tau_t)\mathscr{P}_t. \tag{E.17}$$

E.4. Aggregation and equilibrium

Bond-market clearing requires Ricardian bond holdings to equal government debt,

$$B_{g,t} = (1-\lambda)b_{R,t}. \tag{E.18}$$

Because the Ricardian family owns all installed capital, capital-market clearing is

$$K_{t+1} = (1-\lambda)k_{R,t}. \tag{E.19}$$

Aggregate consumption and the employment rate are, respectively,

$$C_t = \lambda C_{H,t} + (1-\lambda)C_{R,t}, \tag{E.20}$$

$$N_t = \lambda(1 - u_{H,t}) + (1-\lambda)(1 - u_{R,t}), \tag{E.21}$$

and the aggregate resource constraint is

$$\begin{aligned}Y_t = {} & C_t + I_t + G_t + \lambda\iota_H V_{H,t} + (1-\lambda)\iota_R V_{R,t} \\ & + \mathscr{A}_t^p + \mathscr{A}_t^K + \Phi.\end{aligned} \tag{E.22}$$

A competitive equilibrium consists of sequences of household allocations, labor-market flows, production inputs, capital, asset holdings, prices, and policy variables such that households and firms optimize. Matching, free-entry, wage-setting, and price-setting conditions hold. Fiscal and monetary policy satisfy the government budget constraint and the Taylor rule. The labor, capital, bond, and goods markets clear.

E.5. Solution method

The model is linearized around the deterministic steady state and solved with a standard first-order perturbation method (Schmitt-Grohé and Uribe, 2004).

Appendix Table F.1: Calibration: All Parameters

Parameter	Value	Description	Source or target
Panel A: Households			
λ	0.75	Hand-to-mouth population share	College-attainment split
β	0.99	Quarterly discount factor	Smets and Wouters (2007)
σ	1	Logarithmic GHH utility index	Standard value
ξ_H, ξ_R	1/3	Inverse intensive-margin Frisch elasticities	Chetty et al. (2011)
n_H, n_R	1.5, 1	Steady-state hours conditional on employment	Baseline targets
e_H, e_R	1, 2	Labor efficiency units	Normalization
Panel B: Labor markets			
u_H, u_R	0.060, 0.025	Unemployment rates	Current Population Survey
λ_H, λ_R	0.055, 0.022	Quarterly separation probabilities	Wolcott (2021); Dolado et al. (2021)
q_H^f, q_R^f	0.312, 0.710	Vacancy-filling targets	Dolado et al. (2021) (ordering, q_H^f); Den Haan et al. (2000) (q_R^f)
ι_H, ι_R	0.02, 0.08	Vacancy-posting costs	Calibration; Blatter et al. (2012)
Panel C: Production and capital			
a	0.60	Low-skill weight in outer CES nest	Baseline calibration
ζ	0.401	Outer-nest CES parameter	Krusell et al. (2000)
ρ	−0.495	Inner-nest CES parameter	Krusell et al. (2000)
ω_K	0.35	Capital weight in inner CES nest	Dolado et al. (2021)
η	3	Elasticity across differentiated goods	De Loecker et al. (2020)
ϕ_p	50	Price-adjustment cost	Calibration; L'Huillier et al. (2024)
$\delta(1)$	0.0177	Quarterly steady-state depreciation rate	Matches 6.9 percent annually; Cooper and Haltiwanger (2006)
δ_2	3.5	Curvature of utilization depreciation	Baseline calibration
ϕ_K	10	Capital-growth adjustment cost	Matches Basu et al. (2021)
Panel D: Fiscal policy and distribution			
τ	0.25	Steady-state tax rate	Trabandt and Uhlig (2011)
$\bar{B}_g/(4\bar{Y})$	0.20	Debt relative to annual output	Baseline target
ρ_G, ρ_R	0.85, 0.85	Spending and policy-rate persistence	External calibration
χ	0.10	Population share assigned ownership income	Top-decile definition
s^{top}	0.36	Baseline top-10-percent income share	Sample median

Notes: Rates and probabilities are quarterly unless otherwise noted. The internally determined parameters are $\alpha_R = 1.314$, $\alpha_H = 0.727$, $\psi_R = 0.259$, $\psi_H = 0.237$, $\delta_1 = 0.00795$, $\delta_0 = 0.00977$, and $\Phi = 0.131$. The capital stock is determined by no arbitrage and equals $K = 1.742$.

F Calibration details

This appendix documents the calibration of the baseline steady state in Section 5. Appendix Table F.1 reports all parameters set from external evidence, normalizations, and steady-state targets, and its notes list the parameters that are determined internally.

Households. We set the share of hand-to-mouth, low-skill households to $\lambda = 0.75$, so that Ricardian, high-skill households account for the remaining quarter of the population. This division is consistent with interpreting the two groups by college attainment near the beginning of our empirical sample. The quarterly discount factor is $\beta = 0.99$, implying a steady-state annual real interest rate of approximately 4 percent, as in Smets and Wouters (2007). We set the curvature of the GHH utility

index to $\sigma = 1$, so that utility takes the logarithmic limiting form, and the inverse Frisch elasticities to $\xi_H = \xi_R = 1/3$. The latter choice implies an intensive-margin Frisch elasticity of three.[3]

We normalize low-skill efficiency to $e_H = 1$ and Ricardian hours to $n_R = 1$, and set $e_R = 2$ and $n_H = 1.5$. The labor-disutility weights ψ_R and ψ_H are then chosen so that the household first-order conditions reproduce these hours. In the resulting steady state, the skill premium, defined as the ratio of high-skill to low-skill hourly earnings, is 1.91, broadly consistent with existing empirical estimates.

Labor markets. We target unemployment rates of $u_R = 0.025$ and $u_H = 0.060$, guided by education-specific unemployment in the Current Population Survey (CPS, U.S. Bureau of Labor Statistics, 2025). We set quarterly separation probabilities exogenously at $\lambda_R = 0.022$ and $\lambda_H = 0.055$. The higher separation risk of less-educated workers is documented by Wolcott (2021) and Cairó and Cajner (2018). The magnitudes are comparable to the quarterly separation rates used by Dolado et al. (2021).

For vacancy filling, we target a high-skill probability of $q_R^f = 0.710$, matching the aggregate quarterly firm-matching probability estimated by Den Haan et al. (2000), whose matching technology we also adopt in Equation (4.7). We set the low-skill target to $q_H^f = 0.312$, whose ordering relative to q_R^f is consistent with the segmented-matching calibration of Dolado et al. (2021). Given the unemployment targets and exogenous separation rates, steady-state flows imply worker job-finding probabilities of 0.462 and 0.463 in the high- and low-skill markets, respectively, so the low-skill target implies a job-finding probability close to the high-skill market's. Thus, differences in steady-state unemployment by education primarily reflect differences in separation rather than job-finding rates, consistent with the finding of Cairó and Cajner (2018) that more- and less-educated workers have similar job-finding rates but markedly different separation rates.

We normalize the low-skill vacancy-posting cost to $\iota_H = 0.02$ and set $\iota_R = 0.08$. This four-to-one ratio implies more resource-intensive skilled recruitment, consistent with the costly skilled hiring studied by Blatter et al. (2012). Conditional on the unemployment and vacancy-filling targets, the two matching-function curvature parameters are determined from the matching technologies, and steady-state wages are determined by free entry. The resulting values are $\alpha_R = 1.314$ and $\alpha_H = 0.727$.

Production technology and capital. The capital–skill complementarity parameters are $\zeta = 0.401$ and $\rho = -0.495$, based on the estimates in Krusell et al. (2000). They imply an elasticity of substitution of 1.67 between low-skill labor and the capital–high-skill composite and an elasticity of 0.67 between capital and high-skill labor.[4] Thus, an increase in capital is more closely associated with high-skill than with low-skill labor. The capital weight in the inner nest is $\omega_K = 0.35$, following the calibration

[3] This is a baseline calibration choice. For comparison, Chetty et al. (2011) recommend an intensive-margin Frisch elasticity of 0.5 based on microeconomic evidence.

[4] An elasticity above one means that low-skill labor and the composite are gross substitutes, while an elasticity below one means that capital and high-skill labor are gross complements.

in Dolado et al. (2021). We set $\eta = 3$, which implies a steady-state gross markup of 1.5. This relatively high markup is consistent with the evidence on market power in De Loecker et al. (2020). In the baseline economy, we set the low-skill weight in the outer production nest to $a = 0.60$. Together with the remaining production and labor-market parameters, this yields an aggregate labor share of 0.61, broadly consistent with existing empirical estimates. We set the Rotemberg price-adjustment parameter to $\phi_p = 50$, which, together with $\eta = 3$, implies a linearized Phillips-curve slope of 0.04.[5]

The quarterly steady-state depreciation rate is $\delta(1) = 0.0177$, matching the 6.9 percent annual rate used by Cooper and Haltiwanger (2006) for manufacturing capital. Steady-state utilization and the price of installed capital are normalized to one. We set the curvature of the utilization schedule to $\delta_2 = 3.5$. The utilization first-order condition and the depreciation target then imply $\delta_1 = 0.00795$ and $\delta_0 = 0.00977$. Given these values, the capital stock is determined by the steady-state no-arbitrage condition between bonds and installed-capital claims. The capital adjustment cost parameter is $\phi_K = 10$, matching the value used by Basu et al. (2021) for quadratic capital-stock adjustment costs.

Government policies. The steady-state tax rate is $\tau = 0.25$, close to the average U.S. labor-income tax rate reported by Trabandt and Uhlig (2011). Government debt is set to 20 percent of annual output, which implies that roughly 29 percent of the Ricardian family's wealth is held as liquid bonds.[6] Steady-state government purchases are then determined residually by the government budget constraint, and the calibration implies a government-purchases-to-GDP ratio of about 20 percent. The government purchases shock follows the quarterly AR(1) process in Equation (4.10), with $\rho_G = 0.85$, and we set the interest-rate smoothing coefficient in the monetary rule to $\rho_R = 0.85$.

Income concentration. For the measurement of income concentration, the owner group has population mass $\chi = 0.10$. As described in Section 4.1.3, this group is assigned the Ricardian family's financial-claim and profit income in the measured cross section. We choose the fixed operating cost Φ so that the baseline top 10% income share is 36%, close to the sample median of 36.29%. The resulting value is $\Phi = 0.131$. The empirical target is a pre-tax income share including realized capital gains. The model measure uses disposable labor income, after-tax dividends, and capital income net of depreciation, as defined in Equation (4.6).

G Alternative sources of income inequality

We examine three alternative ways of changing income inequality: a zero-sum lump-sum transfer between the two household families, a change in the Ricardian population share, and a change in the elasticity of substitution across differentiated goods. These exercises retain the baseline production weight $a = 0.6$. In each case, we preserve the baseline conditional steady-state elasticity

[5] This is a modest slope, of the same order as estimates in quantitative New Keynesian models (L'Huillier et al., 2024).

[6] If bonds are interpreted as liquid wealth and capital as illiquid wealth, as is common in the HANK literature, this ratio was close to 20 percent in the late 1980s according to the Survey of Consumer Finances.

Appendix Table G.1: Fiscal multipliers under alternative sources of income inequality

	Top 10% income share	Year-2 multiplier	
	(%)	Output	Employment
Baseline	36.0	1.65	1.04
Production shift ($a = 0.5$)	40.7	0.76	1.24
Zero-sum lump-sum transfer	40.7	1.60	1.05
Smaller Ricardian population	40.7	0.72	1.66
Lower markup ($\eta = 3.5$)	33.8	0.94	−0.19

Notes: The top 10% income share refers to total income. Multipliers use the average response in the second year relative to the average response of government purchases in that year. They are not cumulative ratios. Employment is the aggregate headcount employment rate, and its multiplier uses the spending response relative to steady-state output. All shocks have the same impact $\Delta G/\bar{Y}$. The conditional vacancy–rental-rate elasticities retain their baseline values in both labor markets. The Phillips-curve slope is held fixed in the $\eta = 3.5$ exercise.

of vacancies with respect to the rental rate in both labor markets and use spending shocks with the same impact change in government purchases relative to steady-state output. Appendix Table G.1 summarizes the results, with the main production-shift experiment included for comparison.

G.1. Zero-sum lump-sum transfers

We levy a fixed lump-sum tax on the hand-to-mouth family and transfer the proceeds to the Ricardian family. The transfer is chosen to raise the top 10% income share from 36.0% to 40.7%, matching the higher-inequality economy in the main experiment. This redistribution leaves the production structure and household population shares unchanged. Steady-state output, capital, wages, unemployment, and labor-market tightness therefore remain at their baseline values.

Fiscal multipliers are little changed: the year-2 output multiplier is 1.60, compared with the baseline 1.65, and the employment multiplier is 1.05, compared with 1.04. Unlike the redistribution of incremental income in Section 6.1, this transfer is fixed and does not redirect the additional income generated by a spending shock. The Ricardian family's share of the impact income gain remains close to 36%. Raising measured inequality through a redistribution of income levels thus has little effect on fiscal propagation when the underlying structure of the economy is unchanged.

G.2. The Ricardian population share

We next reduce the Ricardian population share, $1 - \lambda$, from 25.0% to 14.3%, again raising the top 10% income share to 40.7%. Relative to the baseline, the year-2 output multiplier falls to 0.72, while the employment multiplier rises to 1.66. This reproduces the qualitative ordering of the main production-shift exercise.

G.3. The elasticity across differentiated goods

Finally, we increase η from 3 to 3.5, lowering the steady-state gross markup from 1.50 to 1.40. Unlike the preceding exercises, this change reduces the top 10% income share, from 36.0% to 33.8%. We

Appendix Table H.1: Multipliers Under an Alternative Labor-Market Assumption

	Output		Employment	
	Year 1	Year 2	Year 1	Year 2
Baseline	3.111	1.650	1.738	1.040
Matched vacancy-rental elasticity, $a = 0.50$ (maintained)	2.459	0.759	1.949	1.238
Fixed wage-rental ratio, $a = 0.50$	1.853	0.513	2.020	1.219

Notes: All cases use the jointly calibrated dynamic parameters in Table 7 and a government-spending shock scaled to the same impact change in G/Y^{ss}.

adjust the price-adjustment parameter ϕ_p to keep the Phillips-curve slope, $(\eta-1)/\phi_p$, unchanged.

Relative to the baseline, both fiscal multipliers fall: the year-2 output multiplier is 0.94 and the employment multiplier is −0.19. The output–inequality ordering is therefore the opposite of the main experiment: lower total-income concentration is associated with a lower output multiplier. Labor-income and total-income concentration also move in opposite directions. The top 10% share of total labor earnings rises slightly, from 12.22% to 12.34%, even as its share of total income falls. Thus, this experiment lowers total-income inequality without reducing labor-income concentration.

H The labor-market assumption and its robustness

This appendix asks how much of the multiplier results in Section 6 depend on the labor-market assumption used in Section 6.5 to construct the segmented labor markets of the high-inequality economy. The maintained assumption holds the conditional local elasticity of vacancy creation with respect to the rental price at its baseline value in each segment,

$$\mathcal{E}^{0}_{V_i,r_i} = \frac{A_i}{(1-\omega_i)[1-(\bar{q}_i^{f})^{\alpha_i}]}, \qquad i \in \{H,R\}, \tag{H.1}$$

where $A_i = 1-\beta(1-\lambda_i)$ and $\omega_i = \bar{w}_i/\bar{r}_i$. This pins down the wage-rental ratio and vacancy-filling probability while leaving wages, vacancies, unemployment, and hours endogenous after a shock. The elasticity describes a local response holding wages, hours, available searchers, and expected continuation values fixed, rather than the full general-equilibrium response of vacancies to a shock. It is not the only assumption consistent with the alternative steady state, so we ask whether the ordering of the two multipliers survives under an alternative assumption about the same two labor markets, holding the jointly calibrated dynamic parameters fixed throughout.

The alternative holds the steady-state wage-rental ratios $\bar{w}_i/\bar{r}_i$ at their baseline values in both segments, leaving unemployment, tightness, and vacancy-filling probabilities to adjust. This replaces Equation (H.1) while retaining the remaining baseline parameter values.

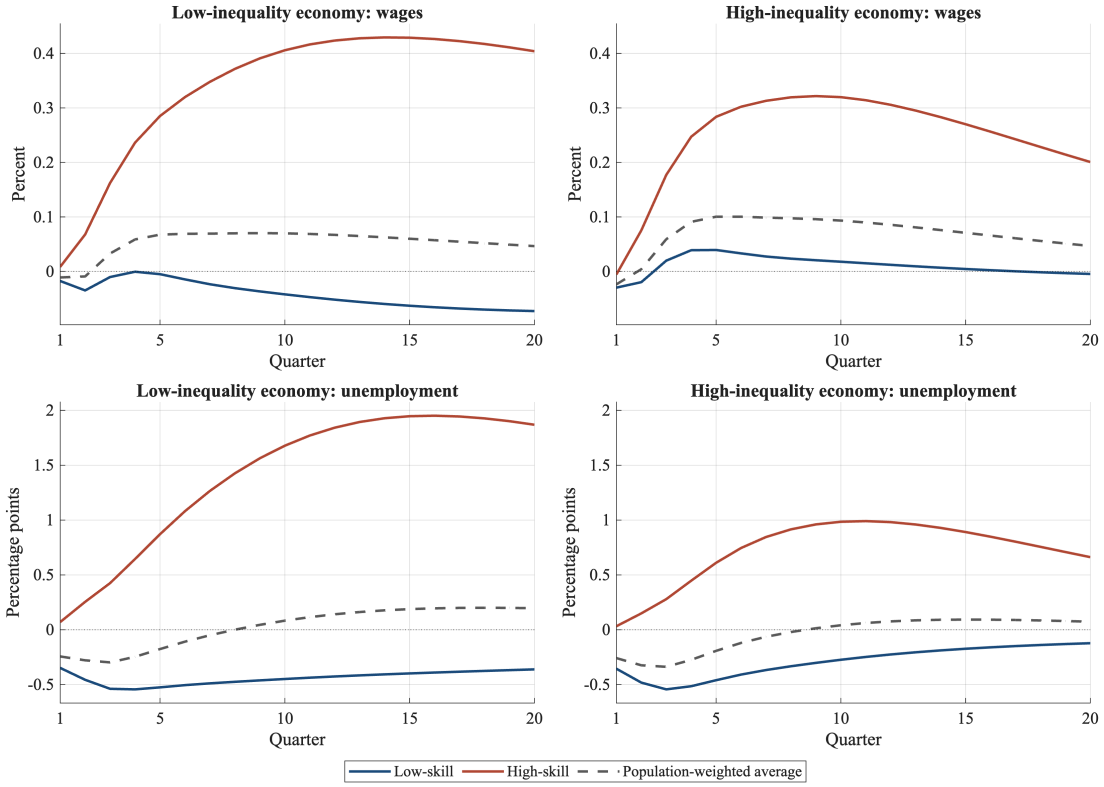
Appendix Table H.1 reports the multipliers for the first two years under each assumption. The central ordering is robust: both assumptions produce a lower year-2 output multiplier and a not-lower year-2 employment multiplier in the high-inequality economy than in the baseline. If

anything, fixing the wage-rental ratios shows a larger decline in the output multiplier, to 0.51 rather than 0.76. Because this alternative does not target any particular local elasticity, this indicates that the central result is not an artifact of matching the vacancy-rental elasticity specifically.

We adopt the elasticity-matching assumption as the maintained specification because it holds fixed the local sensitivity of vacancy creation to the rental price in each segment—the main margin through which the labor market can independently amplify or dampen the effect of the production shift—while leaving wages, unemployment, and hours fully endogenous.[7]

I Labor-market responses by skill segment

Appendix Figure I.1 reports the wage and unemployment responses separately for the low-skill and high-skill segments in each economy.



Appendix Figure I.1: Wage and Unemployment Responses by Labor-Market Segment

Notes: The top row shows wages and the bottom row shows unemployment, with the low- and high-inequality economies ($a = 0.60$ and $a = 0.50$) shown separately in the left and right columns. Each panel plots the low-skill and high-skill segments together with their population-weighted average (λ and $1-\lambda$ weights). Wages are percentage deviations from each economy's own steady state. Unemployment rates are percentage-point deviations. The shocks produce the same impact change in G/Y^{ss} across economies.

[7]This assumption equalizes only one specific margin, the local sensitivity of vacancy creation to the rental price, across the two economies. It does not equalize the underlying job-surplus and matching conditions that generate that sensitivity, nor does it hold fixed hours or the general-equilibrium response of wages. The remaining difference between the two economies' multipliers therefore reflects both the production-incidence channel of Section 6.4 and whatever labor-market differences remain after this one margin has been equalized.

Appendix Table J.1: Decomposition of the Difference in the Output-per-Worker Response

	Year 1	Year 2
Output per worker, $\hat{Y} - \hat{N}$	−0.19	−0.24
– Capital services	+0.01	0.00
– Hours per worker	−0.06	+0.15
– Composition of employment	−0.14	−0.38

Notes: Entries are the high-inequality ($a = 0.50$) minus the low-inequality ($a = 0.60$) response of each term in Equation (J.1), per percentage point of the top 10% income share and per unit of the year-j change in G/Y^{ss}. The three terms sum to the first row.

J Output-per-worker decomposition

This appendix describes the decomposition of the year-2 response of log output per worker used in Section 6.5. Production in Equation (4.9) has constant returns to capital services and the two effective labor inputs, so to first order the log response of output is $\hat{Y} = s_K\hat{K} + s_H\hat{L}_H + s_R\hat{L}_R$, where hats denote log deviations from the economy's own steady state and s_K, s_H, and s_R are the output elasticities of capital services and of low- and high-skill labor. Effective labor in Equation (4.8) is the product of the employment rate and hours per worker, with population shares and efficiency units fixed, so that $\hat{L}_i = \hat{e}_i + \hat{h}_i$, where $\hat{e}_i$ is the log response of the employment rate of group i and $\hat{h}_i$ that of hours per worker. Headcount employment N in Equation (E.21) satisfies $\hat{N} = \sum_i \theta_i \hat{e}_i$, where θ_i is the share of group i in the employed. Subtracting $\hat{N}$ from $\hat{Y}$ gives

$$\hat{Y} - \hat{N} = \underbrace{s_K\hat{K}}_{\text{capital services}} + \underbrace{\sum_i s_i \hat{h}_i}_{\text{hours}} + \underbrace{\sum_i (s_i - \theta_i)\,\hat{e}_i}_{\text{composition}}. \tag{J.1}$$

Capital services differ little across the two economies. Hours matter because the same workers produce more when they work more. The composition term asks whether employment changes add more to output than to headcount: a group with $s_i < \theta_i$ adds more workers than output.

The responses are log changes in year-2 averages divided by the year-2 change in G/Y^{ss}. The weights s_i are output elasticities, computed from factor payments relative to output, and θ_i are employment shares, both at each economy's steady state. Appendix Table J.1 reports the difference between the high- and low-inequality economies per percentage point of the top 10% income share. The three terms sum to the difference in the response of output per worker. Composition is the remainder after capital services and hours. To first order it equals the last term of Equation (J.1) (−0.34 in year 2), and a second-order term adds −0.05. The year-2 hours contribution comes mainly from low-skill workers (+0.13, and high-skill workers contribute +0.02). The model does not reproduce the hours pattern in the data: the unweighted average of the two groups' hours responses is lower in the high-inequality economy, whereas manufacturing hours respond by more in high-inequality states (Appendix B.2). The positive hours contribution reflects the output elasticities that

weight the two groups' responses.

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