Dissecting the Asymmetric Impact of Monetary and Fiscal Policy on Inequality Over the Business Cycle *

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Abstract

Monetary and fiscal policies affect households differently. Using U.S. data from 1980 to 2008 and robustness samples extended through 2019, we document that contractionary policies exacerbate consumption inequality, with a stronger effect during recessions. However, no such asymmetry emerges for earnings inequality. To rationalize this, we propose a New Keynesian model with one key novel feature: loss aversion of households which increases with income. During recessions, the greater loss aversion of high-income households leads them to smooth their consumption by supplying more labor, shifting the burden of contractionary policy onto low-income households and exacerbating consumption inequality. Quantitatively, our model accounts for 50-80% of the asymmetric consumption inequality responses observed in the data.

Keywords: Monetary Policy, Fiscal Policy, Inequality, Loss Aversion

JEL Classification: E32; E52; E62; E21

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1 Introduction

The distributional effects of monetary and fiscal policies have emerged as a critical consideration in macroeconomic policy design.[1] While extensive research documents how these policies affect inequality on average, their distributional impacts across the business cycle are less well understood. This is particularly relevant as policymakers increasingly weigh the equity implications of their decisions alongside traditional macroeconomic objectives. This paper thus aims to study, both empirically and theoretically, how monetary and fiscal policies affect inequality over the business cycles.

We begin by empirically documenting the distributional impact of monetary and fiscal policies using Consumer Expenditure Survey (CEX) data. Our empirical findings are as follows. First, we show that contractionary policy—both fiscal and monetary—tends to exacerbate consumption inequality. Second, we find an asymmetric effect of policies on consumption inequality over the business cycle, with a stronger impact observed during recessions. Third, there is no significant asymmetry observed in the effect of policies on earnings inequality. This suggests that consumption inequality dynamics are not solely driven by changes in earnings.[2]

The second part of our paper proposes a tractable New Keynesian model with heterogeneous households capable of reconciling all our empirical findings. Specifically, we extend an otherwise standard framework by incorporating a single novel feature: loss-averse households, with the degree of loss aversion depending on household income, a feature grounded in recent empirical and experimental findings (see e.g., Gächter et al., 2022; Boyce et al., 2016; Blake et al., 2021). A crucial advantage of our framework is its tractability, allowing us to demonstrate analytically how the model generates state-dependent consumption inequality movement, while no such pattern is present in income. Quantitatively, our framework suggests that the model accounts for over 80% of the short-term and about 50% of the medium-term asymmetric responses of consumption inequality to policy changes. Furthermore, our framework also predicts that no such asymmetry exists for the responses of earnings inequality.

[1] For instance, see the discussion on the impact of monetary policy on inequality between Federal Reserve Chairman Jerome Powell and Senator Elizabeth Warren (transcript), or Treasury Secretary Janet Yellen's statement to the U.S. Chamber of Commerce regarding the impact of fiscal policy on inequality (https://home.treasury.gov/news/press-releases/jy0182).

[2] The first finding aligns with the existing literature (see e.g., Coibion et al., 2017; Anderson et al., 2016, for monetary policy and fiscal policy respectively), while, to the best of our knowledge, the rest has not been previously documented. Notably, although no existing evidence examines the impact of fiscal policy on earnings inequality across business cycles, Furceri et al. (2018) find mixed evidence on the effects of monetary policy on income inequality.

More concretely, our empirical strategy employs state-dependent local projections to estimate differential effects across business cycle states. We identify monetary policy shocks using Romer and Romer (2004) and fiscal policy shocks using the Blanchard-Perotti identification. We measure inequality using both the Gini coefficient and the 90-10 log percentile ratio, constructed from CEX data spanning 1980Q1-2008Q4. Recessions and expansions are classified based on GDP growth relative to trend. Our appendix provides additional robustness checks, including alternative monetary policy shocks from Aruoba and Drechsel (2024) and the high-frequency shock of Jarociński and Karadi (2020), alternative state classifications, and results for alternative inequality measures; we also extend the inequality series through 2019, and the results are unchanged.

In the theoretical framework, our key innovation is introducing loss aversion into households' preferences, where distress from losses outweighs the satisfaction from equal gains. Moreover, we allow the degree of loss aversion to be positively related to income. This assumption generates state-dependent consumption responses in line with the literature (see e.g., Chetty et al., 2024; Fuster et al., 2021), and drives a wedge between the dynamics of consumption and earnings inequality, as we observe in the data.

Our model reveals two fundamental mechanisms that explain why policies have stronger distributional effects during recessions. The first mechanism is common to all households and operates through the curvature of utility. During downturns, when consumption falls below its reference point, households become more willing to substitute consumption over time, so that a given policy change moves aggregate demand by more; policy is thus more potent in recessions. The second mechanism requires heterogeneity in loss aversion and determines which households bear the adjustment. Because high-income households are more loss-averse, they are the most reluctant to let consumption fall below the reference point, smoothing it by supplying more labor, so the burden of the deeper contraction falls on low-income households, who rely primarily on wages and whose consumption moves more than one-for-one with output. Consumption inequality therefore widens in recessions, while earnings inequality does not.

This second channel operates through wages and income composition. Low-income households rely primarily on wages, while high-income households have more diversified income sources including profits. During recessions, this difference becomes crucial. Contractionary monetary policy that reduces economic activity disproportionately impacts low-income households through declining wages. Government spending has the opposite effect—it tends to boost wages while potentially squeezing profits through higher labor costs, providing greater relative

benefit to low-income households during downturns when wage sensitivity is heightened.

Beyond these positive mechanisms, we examine the policy consequences of ignoring state dependence within the model. We show that a central bank ignoring this asymmetry generates substantial welfare costs through three channels: underestimating the general equilibrium multiplier, overestimating the Phillips curve slope, and ignoring the state dependence of intertemporal substitution under loss aversion. Together, these errors cause the central bank to keep interest rates too high during recessions, fighting inflation too aggressively. This policy mistake not only reduces aggregate welfare but also exacerbates consumption inequality beyond what is optimal, with the burden falling on households least able to bear it. We present these results as an illustration rather than as a quantitative policy prescription.

Finally, we discuss the necessity of introducing loss aversion. While the precautionary motive driven by idiosyncratic income risk across business cycles, as in the standard HANK model, could generate similar effects, it alone is insufficient. When we calibrate a model incorporating only idiosyncratic income risk to match income inequality across business cycles, it fails to fully account for the amplified effects on consumption inequality that we observe quantitatively.

This paper is related to several areas of the literature. Empirically, Santoro et al. (2014) and Weise (1999) have pointed out the more substantial effects of monetary policies during recessions compared to expansions. In the fiscal realm, the works of Auerbach and Gorodnichenko (2012a) and Fazzari et al. (2015) emphasize greater policy effectiveness in economic downturns. In contrast, our paper contributes to the empirical literature by providing a simultaneous analysis of the impacts of fiscal and monetary policy on inequality. In addition, we document direct evidence on the state-dependent nature of these effects. That consumption and earnings inequality behave differently is well documented: Krueger and Perri (2006) find that consumption inequality rose far less than income inequality, and Heathcote et al. (2010) that earnings dispersion is more cyclical than consumption dispersion. These studies characterize long-run trends and unconditional cyclical behavior, whereas we compare the responses of the two measures to identified monetary and fiscal shocks across business-cycle states. Theoretically, our model echoes the notions presented by Galí et al. (2007), Kaplan et al. (2014), and Bilbiie (2025). Additionally, our model links the interaction between inequality and the impact of monetary and fiscal policies, contributing to the literature on the state-dependent effectiveness of policies as shown in Bilbiie et al. (2023). However, we show that by integrating the concept of heterogeneous loss aversion, as empirically evidenced by Gächter et al. (2022) and Boyce et al. (2016), our model is capable of explaining the documented empirical patterns.

The remainder of the paper is organized as follows. Section 2 presents our empirical findings. Section 3 describes the model setup and explains how loss aversion is incorporated into the New Keynesian framework. Section 4 outlines the model's analytical results. Section 5 presents the quantitative results, assesses the model's fit with the empirical evidence, and extends it to include idiosyncratic income risk. Section 6 concludes.

2 Empirical Evidence

In this section, we examine the state-dependent effects of monetary and fiscal policy on consumption and earnings inequality. We begin by detailing how we construct our inequality measures, followed by the estimation of our results. Our findings indicate that the impact of policy on consumption inequality varies across different economic conditions, whereas the impact on earnings inequality does not significantly vary across economic states.

2.1 Methodology and Data

To analyze the impact of monetary and fiscal policy on inequality, we employ the local projection method of Jordà (2005). This approach offers several advantages over traditional VAR methods: it is more robust to model misspecification, accommodates state dependence naturally, and does not impose restrictive dynamic relationships between variables. Moreover, it allows for straightforward inference and flexible estimation of impulse responses at different horizons. Given our focus on policy impacts across economic states, we extend the standard local projection framework to incorporate state dependence, following Santoro et al. (2014).[3]

Estimation. Specifically, we estimate a series of state-dependent local projections:

$$\begin{aligned} x_{t+h} = c^{(h)} &+ F(z_t)\left[\sum_{i=0}^{I}(\beta_i^r)^{(h)}\eta_{t-i} + \sum_{j=1}^{J}(\alpha_j^r)^{(h)}x_{t-j}\right] \\ &+ (1-F(z_t))\left[\sum_{i=0}^{I}(\beta_i^e)^{(h)}\eta_{t-i} + \sum_{j=1}^{J}(\alpha_j^e)^{(h)}x_{t-j}\right] + \varepsilon_{t+h}, h = 0, \\ &\ldots, H. \end{aligned} \tag{1}$$

[3]Besides using the Romer and Romer (2004) shock as in Santoro et al. (2014), we also estimate our results using alternative recent monetary policy instruments from Aruoba and Drechsel (2024) as a robustness check. The results remain unchanged. A more detailed discussion is provided later in this section and in the online Appendix A.1. The online Appendix A.5 further extends the analysis to alternative shock series across identification families.

where x_{t+h} is the response variable at horizon h, we choose J=2,I=20 and H=20 to be consistent with Coibion et al. (2017). The function $F(z_t)$ represents the probability of being in a recession, while η_t denotes our identified policy shocks. The coefficients $(\beta_i^e)^{(h)}$ and $(\beta_i^r)^{(h)}$ measure the average effect of policy interventions during expansions and recessions at horizon h.[4] Standard errors are clustered at the quarterly level to account for potential serial correlation and the rotating panel structure of our consumption data.[5]

Dependent Variables. As our main variables of interest, x, we consider two separate measures of earnings and consumption inequality: the Gini coefficient and the (log) difference between the respective values in the 90th and 10th percentiles. The Gini coefficient is computed as $G = \frac{2}{n}\sum_{i=1}^{n} iy_i - \frac{n+1}{n}$, where n is the number of individuals or households, and y_i is the consumption or income of the i-th individual, ranked from the poorest to the richest.

Information on both consumption and earnings is taken from the Consumer Expenditure Survey (CEX). The CEX data is one of the main data sources for household consumption data in the U.S. administered by the Bureau of Labor Statistics (BLS). The CEX is a monthly rotating panel of approximately 1500–2500 households. We consider quarterly information between 1980Q1 and 2008Q4, using the series constructed by Coibion et al. (2017).[6]

As in Coibion et al. (2017), we distinguish between consumption and expenditures. The category of consumption goods encompasses items like non-durables, various services, and selected durable goods, while explicitly omitting substantial acquisitions like houses and automobiles. In contrast, expenditures extend to encompass major expenses, including mortgage instalments, vehicle acquisitions, healthcare products and services, educational fees, and the cost of textbooks, among other items.

For the measurement of earnings, we consider two variables: labor earnings and total in-

[4] As Gonçalves et al. (2024) show, identification problems may arise when policy shocks influence the economic state. We confirm that our policy shocks do not systematically affect the timing of recessions, ensuring the validity of our state-dependent analysis. Detailed test results are reported in the online Appendix Table A.1.

[5] Specifically, the associated standard errors are computed by associated β as in Driscoll and Kraay (1998). The reason is that CEX data surveys households across rotating cohorts, requiring adjustments for cross-sectional dependence as discussed in Coibion et al. (2017).

[6] For results on samples extending beyond 2008, the inequality series are constructed through 2019Q4 from the CEX public-use microdata following Coibion et al. (2017): the same expenditure categories for the consumption and expenditure series, and the same income definitions for the earnings and total-income series, with the Gini averaged over the Bureau's five imputation draws. The extended series are spliced to the original series at 2009. The sample ends in 2019Q4, before the COVID-19 pandemic.

come. While labor earnings are provided directly from the CEX, total income is obtained by summing labor earnings with financial income, business income, and transfers for each household.

Monetary and Fiscal Policy Shocks. The policy shock variable, η_t, represents monetary and fiscal policy shocks at time t. We estimate Eq. (1) separately for each of these policy disturbances. For monetary shocks, we use the monetary policy shock in Romer and Romer (2004) and Aruoba and Drechsel (2024), with the data extended until 2008 subject to data availability.[7] For the fiscal policy shocks, we employ the Blanchard-Perotti identification as in Ramey and Zubairy (2018).[8]

Expansions and Recessions. Expansions (recessions) are defined using periods of positive (negative) growth of a smooth indicator of economic activity, y_t^s. We follow Granger et al. (1993) and use a logistic transformation of y_t^s to define $z_t = \Delta y_t^s$ as the average of GDP within 7 quarters, and $F(z_t) = \exp(-\tau z_t)(1 + \exp(-\tau z_t))^{-1}$. This way, F(z) measures the likelihood of being in a recession, and τ represents how quickly the economy switches from different states. A smooth state, rather than a discrete recession dummy, also matters for estimation: Eq. (1) interacts every regressor with the state, and recessions cover only about one-sixth of the sample, so a discrete indicator would leave too few recession quarters to identify the recession-side coefficients, whereas the continuous $F(z_t)$ lets every quarter inform both regimes. In our case, we use real GDP as the indicator of economic activity, y_t^s. The data for this variable is taken from the Bureau of Economic Analysis after adjusting the price with the GDP deflator.[9]

2.2 Asymmetric Impact of Monetary and Fiscal Policy on Inequality

Figure 1 and 2 presents our main empirical results. Our focus is on the difference in impact between recessions and expansions. Therefore, we show the results of $\beta_j^r - \beta_j^e$ which captures

[7] The Romer and Romer (2004) shock is the change in the intended federal funds rate at each FOMC meeting, purged of the Federal Reserve's real-time (Greenbook) forecasts of output and inflation, so it captures policy changes not driven by the Fed's own economic outlook. On samples, Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 and all end at 2008/12.

[8] The Blanchard-Perotti approach identifies the fiscal shock using institutional information on tax and transfer elasticities to net out the automatic response of the budget to output; because discretionary spending takes more than a quarter to legislate and implement, its within-quarter movement orthogonal to this cyclical component is treated as exogenous. Details on the construction of both policy shocks can be found in the online Appendix A.1.

[9] In the online Appendix Figure A.8, we compare our 7-quarters average recessions indicator with the 2-quarters average with the NBER recession indicator and both indicators are in line with each other.

the difference in average effects of monetary and fiscal policy on inequality during recessions and expansions at horizons j.[10]

Figure 1 shows the impulse responses, $\beta_j^r - \beta_j^e$, for monetary (upper panels) and fiscal policy (lower panels), on the Gini coefficient (left panels) and the 90-10 percentile difference (right panels). We find that positive (contractionary) monetary policy shocks increase consumption inequality significantly more in recessions than in expansions as shown in the upper panel of Figure 1. This observation is in line with existing theories. Coibion et al. (2017) have shown that contractionary monetary policy exacerbates consumption inequality. Similarly, the lower panel of Figure 1 demonstrates that positive (expansionary) fiscal policy reduces consumption inequality more in recession than in expansion. Our result aligns with the finding of Anderson et al. (2016) who show the effectiveness of government spending in mitigating consumption inequality.

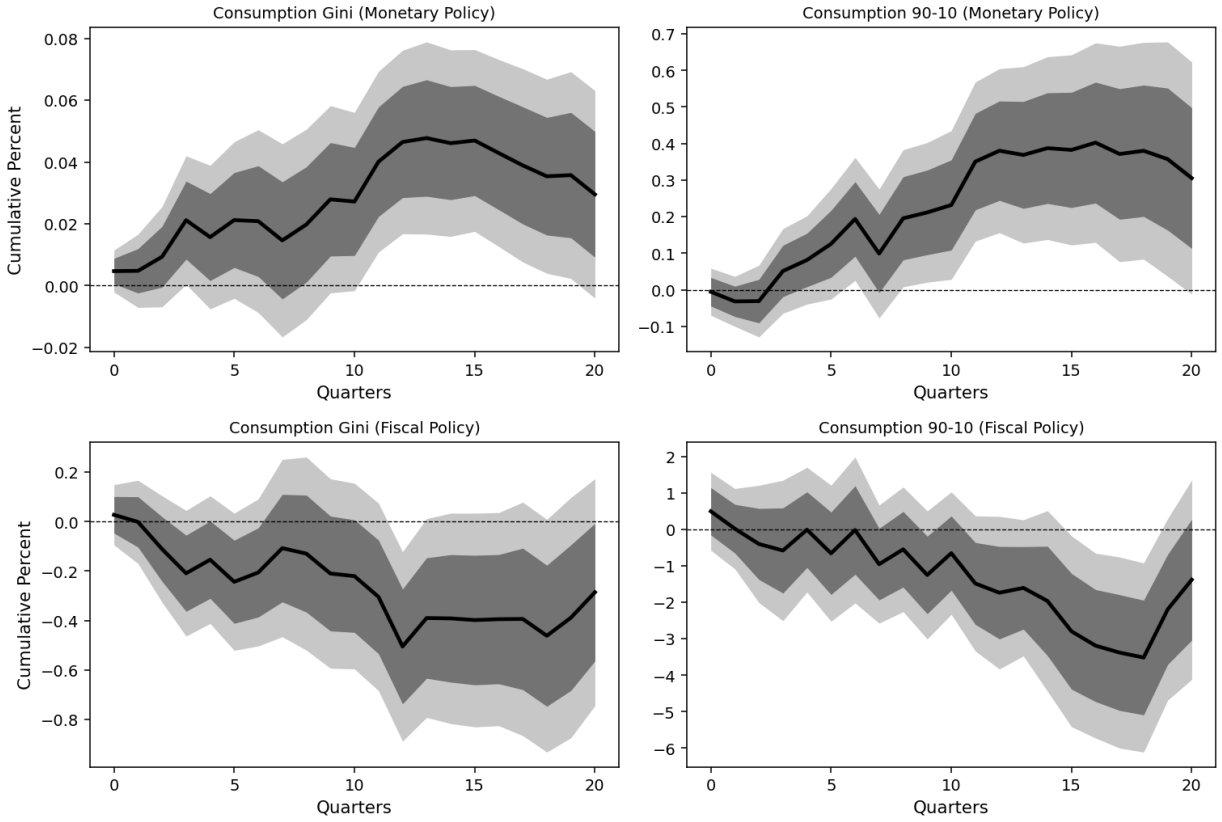
To further investigate the channels through which policy affects consumption inequality, we also examine its impact on earnings inequality. Figure 2 presents the results for earnings inequality. Intuitively, if earnings inequality played a significant role in explaining the state-dependent effects of policy on consumption inequality, we would expect to observe a similar pattern in earnings inequality. However, as shown in Figure 2, the impact of monetary and fiscal policy on earnings inequality remains relatively muted across different economic conditions.

Figures 1 and 2 report the difference between the two state responses, $\beta_j^r - \beta_j^e$. A natural question is whether this difference reflects movement in both states or is driven by one alone. Figure 3 therefore plots the recession response β^r and the expansion response β^e separately, for the baseline monetary and fiscal shocks. For consumption (top row) under a contractionary monetary shock the recession response is positive and builds over the first three years while the expansion response turns negative, so the two states move in opposite directions; under an expansionary fiscal shock the signs flip, the recession response negative and the expansion response positive. In each case the gap between the two states is the object plotted in Figure 1. For earnings (bottom row) the recession and expansion responses overlap around zero under both instruments. The asymmetry is thus a property of the consumption distribution and not of the earnings distribution, and it is present for both monetary and fiscal policy.

A related question is where in the consumption distribution the widening occurs, since the

[10] In the main text we report the Romer and Romer (2004) results and the recession-minus-expansion difference; the online Appendices A.1 and A.2 report the other macroeconomic variables, the alternative inequality measures, and the separate recession and expansion responses. Their direction is the same across the business cycle.

Figure 1: Monetary and Fiscal Policy Impact on Consumption Inequality Across Business Cycle

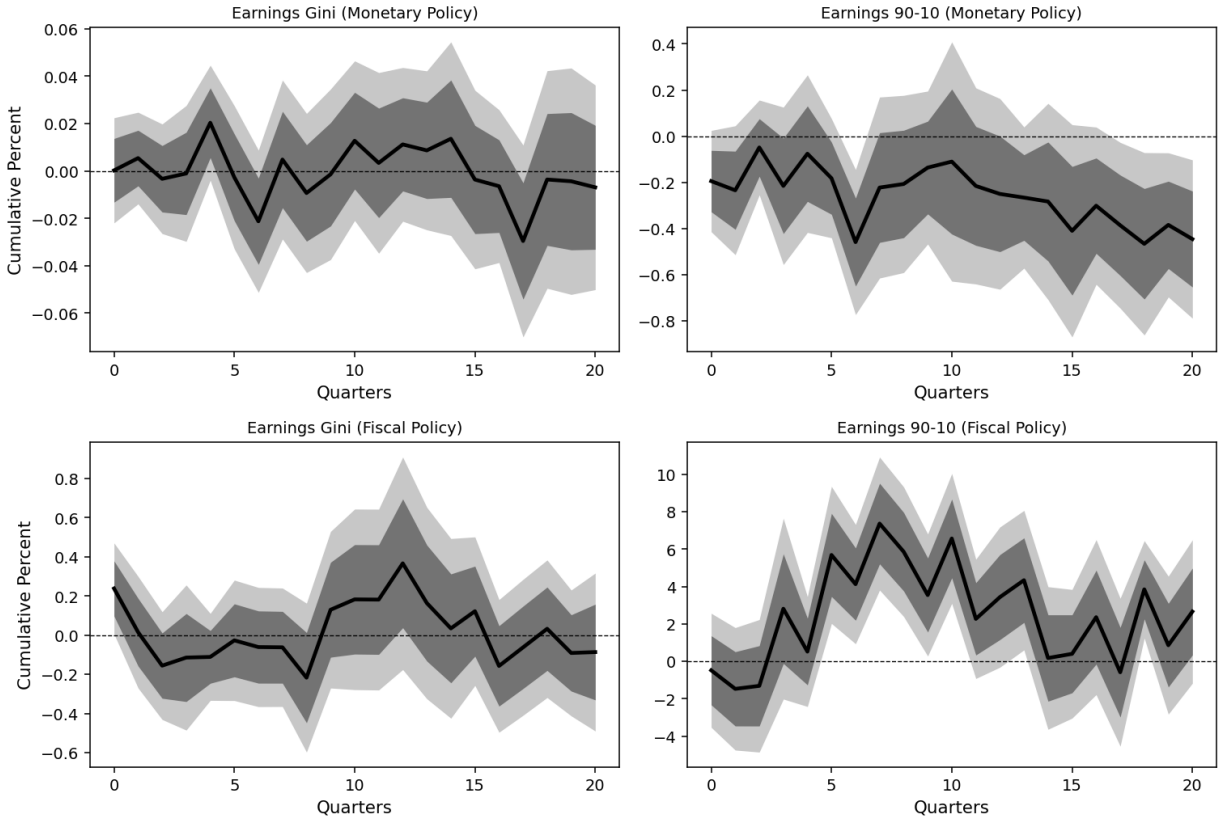


Note: The figure shows the difference in policy impact between recessions and expansions, $\beta_j^r - \beta_j^e$, for a 1% contractionary monetary (top panels) and expansionary fiscal policy shock (bottom panels). It does so for the Gini coefficient (left panels) and 90-10 percentile difference (right panels) of consumption.

90-10 measure is a distance between two percentiles. The online Appendix A.3 therefore re-estimates Eq. (1) separately for the 10th, 50th, and 90th percentiles of log consumption, with the same shock, sample, state, and inference as Figure 1. The 10th percentile is the only leg that declines: its state difference is negative at every horizon and significant at the five-year horizon, whereas neither the median nor the 90th percentile shows a significant decline at any horizon. Repeating the decomposition for earnings shows no corresponding recession-time fall in the 10th percentile of earnings, so the additional consumption adjustment at the bottom of the distribution has no earnings counterpart. This is the incidence pattern of the model of Section 3: in downturns the consumption of the more loss-averse savers stays close to its reference while the loading on low-income consumption steepens, so the consumption gap widens from below even as relative earnings remain stable.

Our baseline shocks follow the studies whose designs we build on: Santoro et al. (2014) estimate state-dependent local projections with the Romer and Romer (2004) monetary shock,

Figure 2: Monetary and Fiscal Policy Impact on Earning Inequality Across Business Cycle

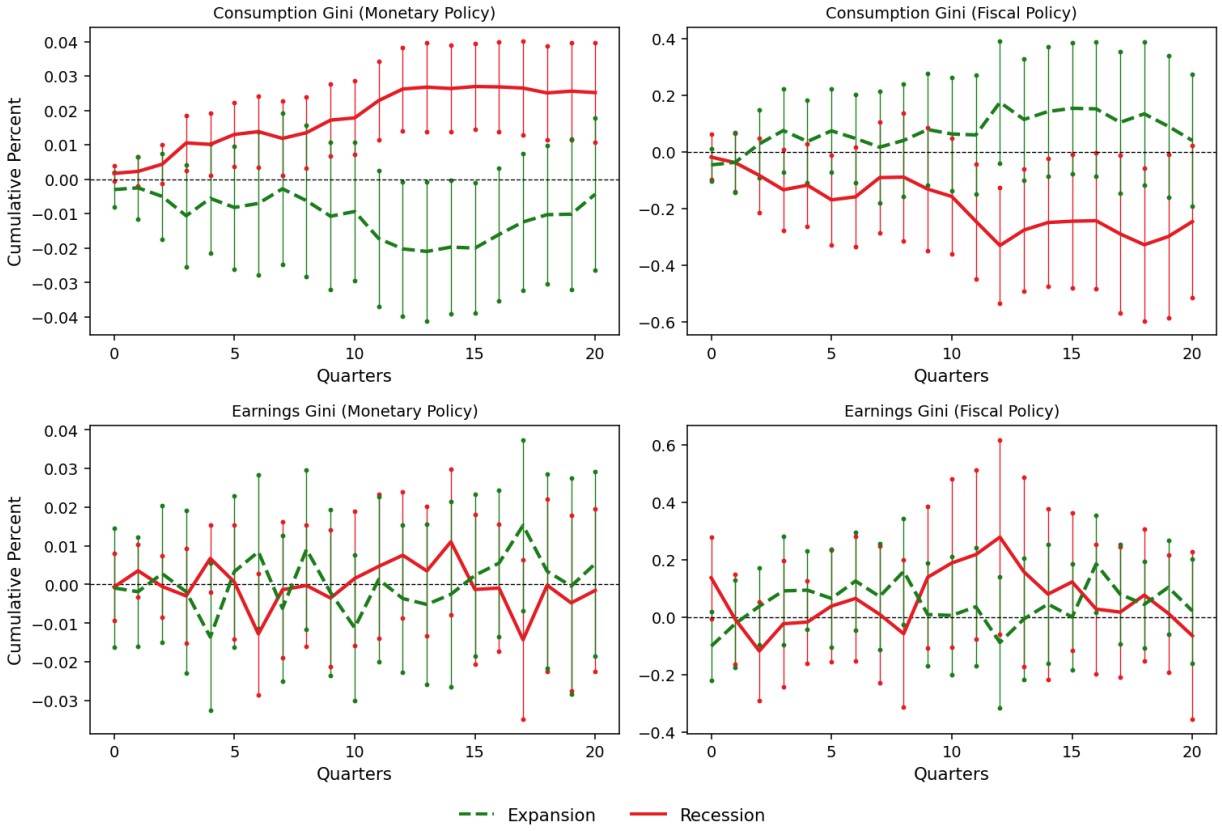


Note: The figure shows the difference in policy impact between recessions and expansions, $\beta_j^r - \beta_j^e$, for a 1% contractionary monetary (top panels) and expansionary fiscal policy shock (bottom panels). It does so for the Gini coefficient (left panels) and 90-10 percentile difference (right panels) of earnings.

and Coibion et al. (2017) use the same shock to document the distributional effects of monetary policy, so the baseline keeps our estimates directly comparable to both, with Aruoba and Drechsel (2024) providing a narrative companion in the same spirit. Three concerns arise with this baseline: the narrative samples end in 2008; the identification of both shocks has been debated, with high-frequency alternatives proposed for monetary policy (e.g., Jarociński and Karadi, 2020; Bauer and Swanson, 2023) and news-based measures for fiscal policy (Ramey and Zubairy, 2018); and the smooth-transition indicator is one of several ways to date recessions. Figure 4 addresses all three at once: we extend the inequality series through 2019, re-estimate the consumption-inequality state difference $\beta^r - \beta^e$ under NBER recession dating,[11] and add two

[11]The discrete NBER indicator leaves too few recession quarters to support the fully interacted lag structure of Eq. (1), so the NBER specification is estimated horizon by horizon with a shorter lag structure, for which Newey-West standard errors are the standard inference; the baseline of Eq. (1) instead pools horizons, and its Driscoll-Kraay intervals are robust to serial correlation and to the correlation across horizons that pooling induces.

Figure 3: Recession versus Expansion Responses of Inequality

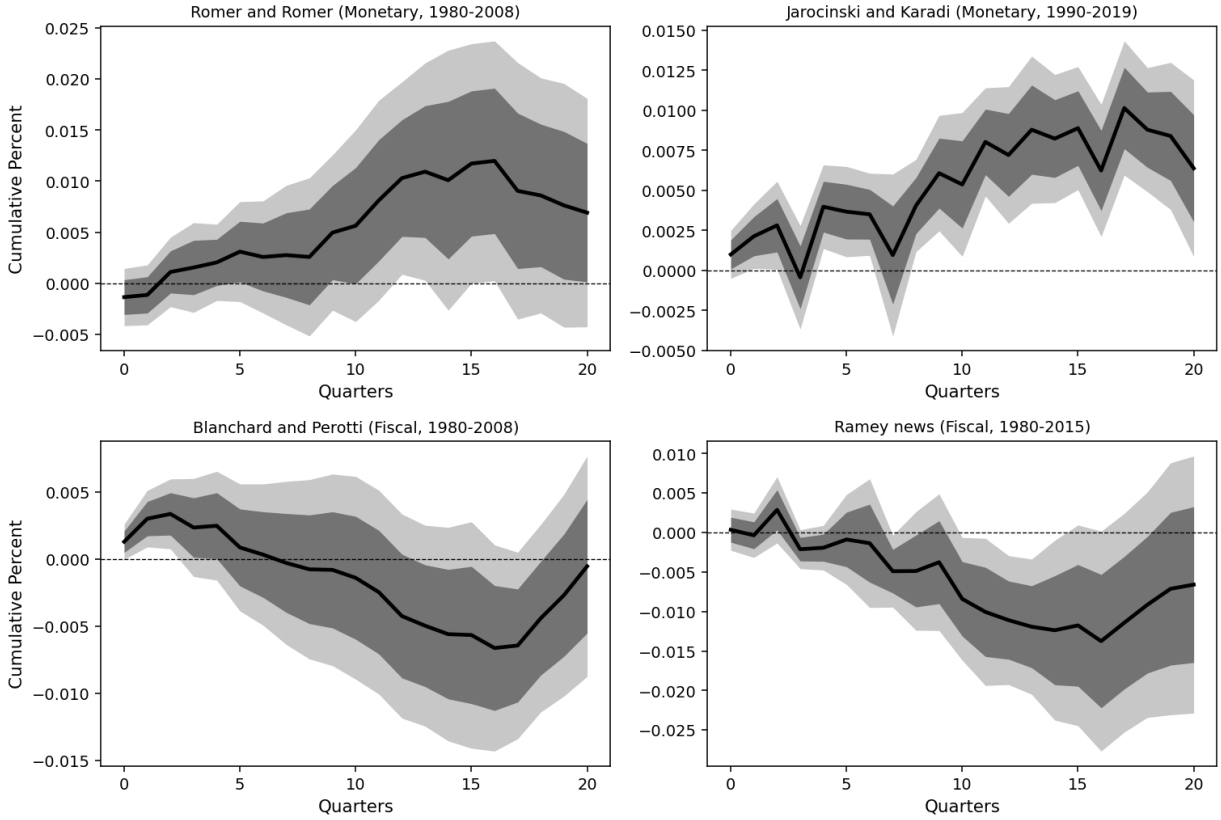


Note: Cumulated responses of the Gini coefficient of consumption (top row) and earnings (bottom row) to a one-standard-deviation contractionary Romer and Romer (2004) monetary shock (left column) and expansionary Blanchard and Perotti (2002) fiscal shock (right column), estimated separately in recessions (β^r, solid red) and expansions (β^e, dashed green) from Eq. (1) on 1980-2008 under the 7-quarter smooth-transition state. Vertical bars are 90% Driscoll-Kraay confidence intervals.

alternative identifications—the high-frequency monetary shock of Jarociński and Karadi (2020) and the Ramey and Zubairy (2018) news fiscal shock[12]—with the asymmetry preserved for both monetary identifications and for the news fiscal shock (under NBER dating the news identification carries the fiscal result). Beyond these, the online Appendix Figures A.19 and A.20 repeat the exercise for the remaining monetary identification families—Gertler and Karadi (2015), Swanson (2021), and Bauer and Swanson (2023), the latter purged of the systematic response of policy to publicly available economic news, the main concern raised about high-frequency

[12]Jarociński and Karadi (2020) identify monetary shocks from the high-frequency co-movement of interest-rate and stock-price surprises in a narrow window around FOMC announcements, separating genuine policy surprises (rates up, equity prices down) from central-bank information shocks (rates and prices move together). The Ramey and Zubairy (2018) news shock measures changes in the expected present value of (largely military) government spending from narrative records, plausibly exogenous to current economic conditions.

Figure 4: Robustness Across Identifications: Consumption-Inequality State Difference under NBER Dating



Note: Cumulated recession-minus-expansion difference $\beta^r - \beta^e$ of the consumption Gini under NBER recession dating (Eq. (A.6)), for a one-standard-deviation contractionary monetary shock (top row: Romer and Romer (2004), 1980-2008; Jarociński and Karadi (2020), 1990-2019) and expansionary fiscal shock (bottom row: Blanchard and Perotti (2002), 1980-2008; Ramey and Zubairy (2018) news, 1980-2015). Dark and light shaded bands are 68% and 90% Newey-West intervals (truncation lag $h + 1$). A positive monetary difference means a recession-timed contraction widens inequality by more; a negative fiscal difference means a recession-timed expansion narrows it by more.

surprises: the consumption asymmetry keeps its sign and state-ordering in every series, and the earnings null holds series by series. Alternative state definitions, the Gonçalves et al. (2024) state-independence test, and per-series estimates are reported in the online Appendices A.3–A.6, and the findings are not driven by particular severe recessions or standard-error specifications.

Given these empirical findings on the asymmetric transmission effects of policies across different economic states, we propose a framework that captures these distinct consumption patterns while preserving fundamental economic mechanisms. In the remainder of this paper, we demonstrate that our approach not only corroborates our empirical findings but also provides an analytical method for examining the interaction between policy and inequality across eco-

nomic cycles. Additionally, we discuss why the conventional model alone may be insufficient to generate the patterns observed in our empirical analysis.

3 Model

We develop a New Keynesian framework with two types of households to explain our empirical findings. Time is discrete and denoted by $t = 0, 1, 2, ...$ The economy consists of a unit mass of households indexed by $i \in [0, 1]$. Each household experiences stochastic transitions between two states $e_{i,t} \in \{H, L\}$ following a Markov process with time-invariant transition matrix $\mathcal{P}$.[13] The stationary distribution of types is denoted by $p(e)$, with a fraction h being high-income savers (H) and $1 - h$ being low-income hand-to-mouth (L). A key innovation in our model is the introduction of heterogeneous loss aversion across household types. This feature is motivated by recent empirical evidence showing that higher-income households exhibit greater loss aversion (e.g., Gächter et al. (2022), Boyce et al. (2016)). As we demonstrate, this heterogeneity in loss aversion, combined with income risk, generates state-dependent policy effects that align with our empirical findings. The framework builds on Santoro et al. (2014): we adopt their gain-loss utility with a reference point and their state-dependent projection design, in which a representative agent's loss aversion makes policy transmit asymmetrically to aggregate variables. Our single departure is to let loss aversion differ across households ($\lambda^H > \lambda^L$) in a two-type economy with unequal profit claims, which is what generates the distributional asymmetry. Setting $\lambda^H = \lambda^L$ in Lemma 2 recovers their representative-agent case exactly.

3.1 Household Preference

To capture the asymmetric behavior of consumption patterns in business cycles and heterogeneous consumption patterns across households during recessions, we assume that household preference features loss aversion and crucially assumes loss aversion is different across households. That is the household will suffer a greater degree of disutility if consumption is less than a reference point compared to the degree of utility it can enjoy if its consumption is more than its reference point and we assume this degree of disutility varies across the household. Specifically, we follow Santoro et al. (2014) and assume the utility for household i at time t as follows:

$$U(C_{i,t}, X_{i,t}) = \alpha V(C_{i,t}) + (1 - \alpha)[G(C_{i,t}, X_{i,t})], \tag{2}$$

[13] The household type is written as a superscript $i \in \{H, L\}$ and the business-cycle regime as $\zeta \in \{e, r\}$; note that e also denotes expansion in regime superscripts such as χ^e.

$$G(C_{i,t}, X_{i,t}) = \begin{cases} \frac{1-\exp(-\theta\delta_{i,t})}{\theta}, & \text{if } \delta_{i,t} > 0 \\ -\lambda(e_{i,t})\frac{1-\exp(\frac{\theta}{\lambda(e_{i,t})}\delta_{i,t})}{\theta}, & \text{otherwise.} \end{cases} ; \quad \theta \geq 0, \quad \lambda(e_{i,t}) \geq 1 \tag{3}$$

where $V(C_{i,t}) = \ln(C_{i,t})$ is a standard log-utility function depends on consumption $C_{i,t}$, $\delta_{i,t} = \ln(C_{i,t}) - \ln(X_{i,t})$ is the "gain-loss" utility term concerning the reference point $X_{i,t}$ which in our case is steady state consumption. This makes all agents in the economy, regardless of their type, go into loss aversion regime when the consumption is below the long run trend.[14], θ is the sensitivity of loss aversion. Crucially, $\lambda(e_{i,t})$ is the coefficient of the loss aversion that is type-dependent, and $(1-\alpha)$ is the degree of loss aversion. Setting $\lambda(e_{i,t}) = 1$ results in the absolute value in the degree of marginal utility being the same between the loss and the gain. Conversely, increasing $\lambda(e_{i,t})$ amplifies the impact of losses relative to gains, making losses weigh more heavily on the household's utility. Noticing that if $\alpha = 1$ the household is not loss averse, our model will converge back to the simple CRRA utility function. As $\theta \to 0$ the gain-loss function becomes linear in $\delta_{i,t}$, as in the piecewise-linear specification of Köbberling and Wakker (2005).

Additionally, households with higher loss aversion face a steeper rise in marginal utility as consumption falls below the reference point, making them more reluctant to let consumption fall below that reference; they smooth it by supplying more labor, so their consumption responds *less*, not more, to a negative shock. We define loss aversion as λ^L for hand-to-mouth type and λ^H for saver type, and assume that high-income (saver) households are more loss-averse than low-income (hand-to-mouth) households, $\lambda^H > \lambda^L \geq 1$. This assumption is supported by behavioral evidence: Blake et al. (2021) show that UK households with low savings are less loss-averse than those with substantial savings, Gächter et al. (2022) find that loss aversion increases with income, and Boyce et al. (2016) find that higher-income groups are more sensitive to losses. The relevant divide is best read as savers versus hand-to-mouth—a distinction in *wealth* rather than a pure income gradient—so the less loss-averse type is the low-wealth hand-to-mouth household (Blake et al., 2021), even though it sits closer to the borrowing constraint. Behaviorally, it implies that hand-to-mouth households reduce consumption less and respond more to income gains during recessions, in line with Chetty et al. (2024) and Fuster et al. (2021).

[14] In Santoro et al. (2014), they assume that reference point is in the habit formation form such that $X_{i,t} = (C_{i,t-1})^{\gamma}$ where γ captures the degree of habit formation; the online Appendix C derives and simulates the habit-reference version and shows the monetary results are preserved. The fiscal block is derived only under the trend reference point.

3.2 Household Problem

Each household i faces a dynamic optimization problem, choosing consumption ($C_{i,t}$), labor supply ($N_{i,t}$), and bond holdings ($B_{i,t}$) to maximize expected lifetime utility:

$$\max_{C_{i,t},N_{i,t},B_{i,t}} E_t \sum_{s=0}^{\infty} \beta^s \left[U(C_{i,t+s}, X_{i,t+s}) - \frac{(N_{i,t+s})^{1+\phi}}{1+\phi} \right], \tag{4}$$

subject to the nominal budget constraint:

$$P_t C_{i,t} + P_t B_{i,t} \leq R_{t-1} B_{i,t-1} + P_t W_t N_{i,t} + P_t Div(e_{i,t}) + P_t T(e_{i,t}), \tag{5}$$

and a borrowing constraint:

$$B_{i,t+1} \geq \underline{B}, \tag{6}$$

where $\beta \in (0,1)$ is the discount factor and $\phi > 0$ governs the Frisch elasticity of labor supply. Income consists of labor earnings ($W_t N_{i,t}$), dividend payments ($Div(e_{i,t})$), and government transfers ($T(e_{i,t})$). Following Bilbiie (2020), dividend income varies by type, with low-income households receiving a smaller share:

$$Div(e_{L,t}) = \frac{\mu^d}{1-h} D_t, \qquad Div(e_{H,t}) = \frac{1-\mu^d}{h} D_t, \tag{7}$$

where D_t is the net profit of the firm and $\mu^d < (1-h)$ captures the proportion of profit of the low income, so that high-income households claim the remaining share $1-\mu^d > h$.[15] Government transfers $T(e_{i,t})$ are type-dependent and used to finance government spending. The real bond $B_{i,t}$ pays nominal interest rate R_t subjected to borrowing constraint $\underline{B}$, and P_t denotes the aggregate price level.

3.3 Firms

Final-goods producers operate under perfect competition, assembling a continuum of intermediate goods with elasticity of substitution $\epsilon > 1$; cost minimization yields the standard demand curves $C_t(z) = (P_t(z)/P_t)^{-\epsilon} C_t$ and the associated price index P_t (online Appendix B). Intermediate firms produce with linear technology $Y_t(z) = A_t N_t(z)$, where the common productivity A_t follows an AR(1) process, and reset prices with probability $1-\Theta$ as in Calvo (1983). Following

[15] This specification bridges TANK models where hand-to-mouth households simply consume their income (e.g., Campbell and Mankiw (1991), Galí et al. (2007)) and HANK models with endogenous wealth distribution (e.g., Kaplan et al. (2014)). As shown by Bilbiie (2020), varying dividend shares can replicate key features of both frameworks.

Bilbiie (2020), a constant revenue subsidy τ_s financed by a lump-sum tax T_t^F on firms removes the steady-state markup, so profits are

$$D_t(z) = (1+\tau_s)\left[\frac{P_t(z)}{P_t}\right] Y_t(z) - W_t N_t(z) - T_t^F, \tag{8}$$

with the wage W_t as marginal cost. Optimal pricing and the subsidy $\tau_s = (\epsilon - 1)^{-1}$ (derivation in the online Appendix B) leave zero net profit in steady state, $D_t = Y_t - W_t N_t = 0$. The subsidy is redistributive: it taxes the holders of firm shares, which matters because dividend claims are concentrated among high-income households.

3.4 Monetary Policy, Fiscal Policy, and Market Clearing Conditions

We assume that the monetary policy maker sets the nominal interest rate $i_t = R_t - 1$ to satisfy the standard Taylor's Rule:

$$i_t = \phi_\pi \pi_t + v_t^m, \tag{9}$$

with $\pi_t = P_t/P_{t-1} - 1$ being the inflation, parameter ϕ_π capturing how monetary policy responds to the inflation, v_t^m being monetary policy shock following an AR(1) process:

$$v_t^m = \rho_m v_{t-1}^m + \epsilon_{mp,t}, \tag{10}$$

where ρ_m is the persistence of monetary policy shock, $\epsilon_{mp,t} \sim N(0, \sigma_m^2)$ is the i.i.d normal distribution.

For fiscal policy, the government needs to satisfy the period-by-period budget constraint such that:

$$P_t T_t + R_t^{-1} B_{t+1}^G = B_t^G + P_t G_t \tag{11}$$

where B^G denote the real bonds issued by the government and $T = hT^H + (1-h)T^L$ is the total uniform tax that is raised to finance the government spending and deficit. Thus, the government spending and taxes follow a simple debt feedback rule:

$$g_t - t_t = (1-\phi_g)\frac{B_{t+1}^G - B^G}{Y}, \tag{12}$$

where Y and B^G are the steady-state values of output and government bond, $g_t = (G_t - G)$, and $t_t = (T_t - T)$ are the value of government spending and tax deviation from the steady state subjected to the deficit to output ratio, $1 - \phi_g$. In addition, government spending follows an AR(1) process:

$$g_t = \rho_g g_{t-1} + \epsilon_{g,t}, \tag{13}$$

where ρ_g is the persistence of government spending shock, $\epsilon_{g,t} \sim N(0, \sigma_g^2)$ is the i.i.d normal distribution.

Equilibrium definition Given the initial price level P_{-1} and government debt B_0^G, an equilibrium is a path for prices, aggregates, and type-specific allocations $\{C_t^i, N_t^i, B_{t+1}^i\}_{i\in\{H,L\}}$ such that households and firms optimize, the policy rules hold, and the goods and bond markets clear: $hC_t^H + (1-h)C_t^L = Y_t - G_t$ and $hB_{t+1}^H + (1-h)B_{t+1}^L = B_{t+1}^G$ (with type switching, h and $1-h$ are the stationary type probabilities).

Asset markets and the zero-liquidity equilibrium The only asset is the one-period government bond, and we study the zero-liquidity equilibrium ($B^G = \underline{B} = 0$): no household carries bonds across periods. High-income (H) households are unconstrained, so their Euler equation holds with equality and prices the bond; low-income (L) households are at the borrowing constraint and consume their period income.[16] Zero net debt turns the debt-feedback rule into period-by-period budget balance, $g_t = t_t$, with the type incidence of the lump-sum tax (t^L versus t^H) entering the analytics. A household that switches from H to L also switches its loss-aversion coefficient from λ^H to λ^L, since λ is attached to the wealth type rather than to the individual household. Types are permanent in the analytics of Section 4, so this matters only in the quantitative model of Section 5, where the risk of switching to the low-λ, low-dividend type acts as the income-risk channel.

4 Analytical Results

To solve the model in closed form, we follow the equilibrium of Pfäuti and Seyrich (2022) and Bilbiie (2025) by solving the model with zero net liquidity such that $B^G = \underline{B} = 0$. Defining the high-income type household's probability to remain in her current type $p(e_{t+1} = H|e_t = H) = s$ and similar for low-income type as $p(e_{t+1} = L|e_t = L) = s_L$. We have $1-h = \frac{1-s}{2-s-s_L}$ as the share of low-income type households as a constant. To fit with empirical findings such that there is no asymmetric effect on wage-earning, we assume the productivity of two types of households is

[16] Consumption inequality in the model therefore operates through the savers' intertemporal (Euler) margin: with no bonds held in equilibrium, desired saving shows up in the real interest rate and in the savers' consumption path rather than in asset holdings. The percentile decomposition of the online Appendix A.3 supports this margin empirically: after a recession-timed contraction, consumption falls at the bottom of the distribution with no matching fall in earnings at the bottom.

the same and thus the only source of income heterogeneity is via the dividend difference.

To understand how policy impacts consumption inequality across business cycles analytically, we focus on the stationary equilibrium such that the probability of remaining in the type is the same such that $P(e_{t+1} = H|e_t = H) = P(e_{t+1} = L|e_t = L) = 1$ and the case when the budget constraint of H household always bind and is indeed a strict inequality for L household. This will nest our model to the tractable TANK model as in Bilbiie (2025).[17] Second, the probability of switching states is the same across periods such that $P(\zeta_{t+1} = e|\zeta_t = e) = P(\zeta_{t+1} = r|\zeta_t = r) = 1$ with $\zeta = \{e, r\}$ is the indicator variable of state where e represents expansion and r represents recession. These assumptions allow us to solve the model as two agents model in different states separately.

Finally, to be consistent with the empirical strategy we adopt in state-dependent local projection where we use the 7-quarters average to determine the economic state, we set the reference point as long-run economic growth, denoted by $X_t = C$ with C being the steady state value of consumption. We can thus interpret the reference point X_t as the long-run economic trend we employ in estimating economic state. During expansionary periods, households' consumption surpasses their reference levels, and the opposite occurs during contractions as the business cycle defined in Santoro et al. (2014).[18] Lowercase letters signify log deviations from the steady state.

4.1 State-Dependence of Key Model Features

In this section, we present the analytical result for household consumption response to the output ratio, IS curve, and NKPC. The result of the demand side (i.e., the IS curve) is obtained by using the first-order condition of household decision and the result for the supply side(i.e., NKPC) is derived from the first-order condition of the firm's problem.

[17] In this case, as shown in Bilbiie (2020), we treat the proportion of the low-income households 1-h as a free parameter. Later, we release this assumption and make the model nested into the type-switching framework of Bilbiie (2020) to show the connection of switching risk with our loss aversion mechanism.

[18] The online Appendix C repeats the simulation with an adaptive (habit) reference point and the monetary asymmetries survive. The empirical recession states are dated by the smooth-transition indicator and by NBER contractions, not by the model's reference point (online Appendix A.6).

State-Dependence Consumption Response Using the labor market condition[19], we obtain the result of the state-dependent nature of consumption response to output:

$$\begin{cases} c^L_{\zeta,t} = \chi^\zeta(y_{\zeta,t} - g_{\zeta,t}) + h(\psi_{g,\zeta} g_t - \psi_{\tau,\zeta} t^L_{\zeta,t}), \\ c^H_{\zeta,t} = \frac{1-(1-h)\chi^\zeta}{h}(y_{\zeta,t} - g_{\zeta,t}) - (1-h)(\psi_{g,\zeta} g_t - \psi_{\tau,\zeta} t^L_{\zeta,t}) \end{cases} \quad \text{if } \zeta = [e, r]. \tag{14}$$

where

$$\chi^e = 1 + \phi\left(1 - \frac{\mu^d}{1-h}\right), \tag{15}$$

$$\chi^r = \frac{(\phi + \sigma^H_r)\chi^e}{(\phi + \sigma^H_r)\chi^e - \underbrace{(\chi^e - 1)(\phi + \sigma^L_r)}_{\text{Indirect Effect}} - \underbrace{h\chi^e(\sigma^H_r - \sigma^L_r)}_{\text{Direct Effect}}}, \tag{16}$$

$$\underbrace{\sigma^i_e = 1 + (1-\alpha)\theta \ \ \forall i}_{\text{Common Effective Curvature in Expansion}} \quad , \quad \underbrace{\sigma^i_r = 1 - (1-\alpha)\theta(\lambda^i)^{-1}}_{\text{Heterogeneous Effective Curvature in Recession}} \quad , \tag{17}$$

$$\psi_{g,e} = \frac{\chi^e}{h}\left(\frac{\phi}{\phi + \sigma^H_e}\right), \psi_{g,r} = \frac{\chi^r}{h}\left(\frac{\phi}{\phi + \sigma^H_r}\right), \tag{18}$$

$$\psi_{\tau,e} = \frac{1}{h}\left(\frac{\phi}{\phi + \sigma^H_e}\right), \psi_{\tau,r} = \frac{1}{h}\frac{\chi^r}{\chi^e}\left(\frac{\phi}{\phi + \sigma^H_r}\right). \tag{19}$$

To understand the features of the model, we first assume there is no government spending (i.e. $g = t = 0$), we begin by analyzing how consumption responds to aggregate conditions in the two-agent framework.

In Eq (15), χ capture the response of low-income consumption to output changes, a general-equilibrium loading rather than a partial-equilibrium marginal propensity to consume out of a transfer. The magnitude of this response is determined by two key factors: the elasticity of labor supply (ϕ) and the relative share of dividend claims ($\frac{\mu^d}{1-h}$). When low-income households receive a smaller share of profits relative to their population proportion, consistent with empirical evidence (e.g.,McKay and Wolf (2023), Bilbiie (2020)), their consumption response to output exceeds unity ($\chi > 1$). This mechanism, highlighted by Bilbiie (2020), arises from the asymmetric exposure to income sources: low-income households rely primarily on wages, while high-income households receive substantial profit income that helps insulate them from output fluctuations. Consequently, when output falls, the decline in wage income disproportionately impacts low-income households, widening consumption inequality. This feature helps explain why contractionary monetary policy, by reducing aggregate demand, increases consumption inequality as documented by Coibion et al. (2017). However, the standard mechanism alone

[19]Detailed derivations are presented in the online Appendix B

cannot account for our empirical finding that consumption inequality responds asymmetrically to output across business cycles. To explain this pattern, we introduce heterogeneous loss aversion across household types. Following Santoro et al. (2014), loss aversion operates through a *state-dependent effective curvature* of utility: below the reference point the coefficient σ^i_ζ, which governs how fast marginal utility rises as consumption falls, is lower than in expansions for both types ($\sigma^i_r < 1 < \sigma^i_e$), raising the elasticity of intertemporal substitution, and it becomes type-dependent. Inspecting σ^i_r, we first establish the ranking of the effective curvatures:

Lemma 1. $\sigma^H_r > \sigma^L_r$

Intuitively, both types' curvature falls in the loss regime, but the decline is smaller for the more loss-averse type, so the high-income household will retain a higher effective curvature—and hence a stronger consumption-restraint motive—than the low-income during recessions. This result follows directly from our assumption that high-income households exhibit greater loss aversion ($\lambda^H > \lambda^L$). Building on this feature, we can show:

Lemma 2. $\chi^r \geq \chi^e$

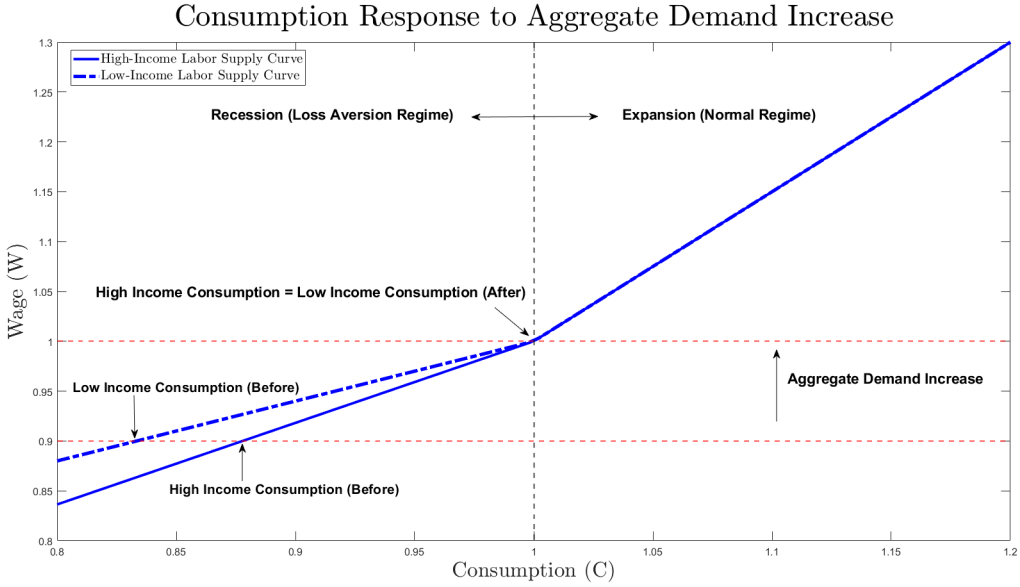
The two lemmas show how the model relates to Santoro et al. (2014). A representative agent carries a single loss-aversion coefficient, so the recession curvature $\sigma_r = 1 - (1-\alpha)\theta\lambda^{-1}$ of Eq. (17) is common across households and loss aversion transmits asymmetrically only to *aggregate* consumption. Here that curvature is type-specific, and the state-dependence of the *distribution* is carried entirely by the wedge

$$\sigma^H_r - \sigma^L_r = (1-\alpha)\theta\left[(\lambda^L)^{-1} - (\lambda^H)^{-1}\right] > 0, \tag{20}$$

which is positive under the assumption $\lambda^H > \lambda^L$. By Lemma 2 the low-income output loading rises in recessions, $\chi^r > \chi^e$, exactly to the extent that (20) is positive, and the consumption gap $\gamma_t = \frac{1-\chi^\zeta}{h} y_t$ inherits this state-dependence through χ^ζ. Setting $\lambda^H = \lambda^L$ closes the wedge and restores $\chi^r = \chi^e$: loss aversion still amplifies transmission in recessions, but symmetrically across types, so the aggregate asymmetry of Santoro et al. survives while the state-dependence of consumption inequality—the fact this paper documents—vanishes. Heterogeneous loss aversion is thus what turns their asymmetric *aggregate* transmission into asymmetric transmission to the *distribution*.

Using Lemma 1 and given that $\chi^e > 1$, a distinctive feature emerges during recessions when consumption falls below reference levels: the response of low-income household consumption

Figure 5: Heterogeneous Labor Supply During Business Cycle



This figure illustrates the labor supply curves for low-income households (blue dashed line) and high-income households (blue solid line) under varying labor market conditions, analyzing the impact of a wage increase from 0.9 to 1 on consumption. The reference consumption level and labor demand are set at 1, with areas to the left representing the recession (loss) regime and areas to the right representing the expansion (gain) regime. Parameter values used in this analysis are provided in Table 1.

becomes even stronger through two channels: a direct effect via heterogeneous effective curvature and an indirect effect through output, which we discuss in detail below.

Direct Effect $h\chi^e(\sigma_r^H - \sigma_r^L)$: This term in equation (16) captures the impact of heterogeneous effective curvature across households during recessions, arising from differences in loss aversion. This consumption-restraint motive during recessions makes the household less likely to consume and prefer to work more compared to expansions resulting in a regime-dependent labor supply curve as in Santoro et al. (2014).

High-income households, being more sensitive to consumption losses, have a stronger motive to maintain their consumption levels. As a result, they are less willing to increase their consumption when wages increase during recessions. In contrast, low-income households, being less loss-averse, are more willing to increase their consumption when facing the same wage increase.

Figure 5 illustrates the labor supply curves for high-income and low-income households. During normal times (right of the reference point), both household types have similar labor supply curves. However, during recessions (left of the reference point), the labor supply curves

diverge due to the differences in loss aversion. Since the labor demand is inelastic in the New Keynesian model and we assume the wage source is the same across the household, when there is the same increase in wage due to an increase in aggregate demand, either due to monetary or fiscal policy, the consumption of the low income will increase more than the consumption of the high income. Our model is consistent with empirical evidence such that households with larger wealth or income are more likely to smooth their consumption and work more compared to the low-income type during the recessions (see e.g. Parker and Vissing-Jørgensen (2009),Blundell et al. (2016)).

This impact will be amplified by χ^e which measures how large the output multiplier effect on consumption as the low-income consumption response is more than one-to-one and depends on the ratio of the high-income household (h). Intuitively, if the low income is more responsive to the output (i.e. higher χ^e), this effect should be larger. Similarly, since this effect stems from the difference in effective curvature between the high and the low income, the effect will be larger when the proportion of the high-income is higher.

By the wedge (20), the *heterogeneity* of loss aversion, rather than its level, generates the state-dependent consumption response: the homogeneous case ($\sigma_r^H = \sigma_r^L$) keeps the common potency channel but produces no distributional asymmetry. This aligns with evidence that low-income households are more responsive to gains in recessions (e.g., Gross et al. (2020), Bilbiie et al. (2022), and Auclert et al. (2020)).

Indirect Effect $(\chi^e - 1)(\phi + \sigma_r^L)$: This mechanism highlights that during recessions, low-income consumption is more sensitive to wage fluctuations than high-income consumption due to differences in income sources. While both household types face wage reductions during output drops, high-income households are less affected because they have a larger share of firm profits to buffer against these drops. The term $(\chi^e - 1)(\phi + \sigma_r^L)$ encapsulates this dynamic, where $(\chi^e - 1)$ measures the extent of this response and the term $(\phi + \sigma_r^L)$ indicates the disparity in output responsiveness as we discuss before that the low-income consumer will consume more and when they consume more, the aggregate consumption will increase.

We now introduce the impact of government spending on output, defined as the sum of total consumption and government expenditure, $y = g + c$. The output, net of government spending, forms the first component of Equation (14). The second component highlights that government spending boosts consumption among low-income groups by increasing overall demand and resulting in higher wages. These higher wages generate higher marginal labor costs and lower firm profits. Since low-income individuals hold a smaller share of these profits compared to

high-income individuals, the net effect of government spending predominantly benefits low-income group consumption. For the third component, the term t_t^L represents the tax amount paid by low-income individuals. This component shows if low-income individuals anticipate higher taxes, they are likely to curtail their spending.

During recessions, the impacts of all three components are magnified. The first component is the same as in the monetary policy discussion such that loss-aversion heterogeneity will lead to a heterogeneous response to output and heterogeneous effective curvature. For the second component, due to loss-averse behavior, the consumption of high-income households is more sensitive to the associated taxes and government spending and thus the high-income will consume relatively less than the low-income when an increase in government spending during recessions. Similarly, loss aversion will also make low-income households more sensitive to tax change and amplify the third component.

IS Curve To obtain the IS curve, we derive the Euler equation via the first-order condition of the optimizing problem of high-income households as shown below.

$$c_{\zeta,t}^H = E_t c_{\zeta,t+1}^H - \frac{1}{\sigma_\zeta^H}(i_t - E_t\pi_{t+1}), \quad \text{if } \zeta = [e, r]. \tag{21}$$

Equation (21) signifies that the savers' intertemporal substitution is state-dependent: in the loss regime their effective curvature is lower ($\sigma_r^H < \sigma_e^H$), so a given change in the real interest rate moves their consumption by more, making monetary policy more powerful during recessions than expansions, as shown in Santoro et al. (2014).

Specifically, incorporating government spending does not alter the consumption Euler equation for high-income savers and they continue to act as Ricardian agents. Consequently, we substitute the state-dependent consumption response of the high-income we obtain from Equation (14) into the consumption Euler equation of the high-income households and we use $c = y - g$ to present the output net of government spending, yielding the IS curve:

$$c_{\zeta,t} = E_t c_{\zeta,t+1} - \psi_{c,\zeta}(1-h)[\psi_{g,\zeta}(E_t g_{t+1} - g_t) - \psi_{\tau,\zeta}(E_t t_{t+1}^L - t_t^L)] - \frac{\psi_{c,\zeta}}{\sigma_\zeta^H}(i_t - E_t\pi_{t+1}), \quad \text{if } \zeta = [e, r], \tag{22}$$

where

$$\psi_{c,e} = \frac{h}{1-(1-h)\chi^e}, \psi_{c,r} = \frac{h}{1-(1-h)\chi^r}. \tag{23}$$

We first investigate the impact of monetary policy on output by looking at the coefficient $\psi_{c,\zeta}$ and we can observe that $\psi_{c,r} > \psi_{c,e}$ indicating a more pronounced policy effect during

recessions. Intuitively, since low-income households are more sensitive to the demand shock than high-income households if the consumption response of the low-income household (i.e., χ^{ζ}) is larger, they are more likely to change their consumption correspondence to the interest rate change. Thus, the consumption response of the low-income household amplifies the impact of the interest rate on output. Additionally, the impact of the monetary policy will be larger during the recession as the common decline in effective curvature raises the sensitivity of consumption to the real interest rate for both households. Hence, the impact of the monetary policy will be larger during the recession period.

In exploring the fiscal policy's effect on output, we note that $\psi_{g,e} = \frac{\chi^e}{h}\left(\frac{\phi}{\phi+\sigma_e^H}\right)$ and $\psi_{g,r} = \frac{\chi^r}{h}\left(\frac{\phi}{\phi+\sigma_r^H}\right)$. Coupled with the insights from Lemma 2, we can show that $\psi_{g,r} > \psi_{g,e}$. With the assumption of government spending following a stationary process, defined by $g_t = \rho_g g_t + \epsilon_{g,t}$ and $\rho_g < 1$, the expected difference between future and current government spending $(E_t g_{t+1} - g_t)$ is typically negative under rational expectations. Therefore, we demonstrate that net government spending boosts output more effectively during recessions in line with previous literature (e.g., Auerbach and Gorodnichenko (2013), Ghassibe and Zanetti (2022)).

The effectiveness of government spending in this model is attributed to the existence of two agent types with distinct claims on firm profits. Government spending boosts output and household wages, which reduces firm profits. While all households benefit from the wage increase, high-income households, which hold larger shares of firm profits, face greater financial losses. These households, acting as Ricardian agents, experience a net loss, heightening their sensitivity due to their higher loss aversion. Conversely, low-income households, with smaller profit shares, gain from the increased wages, leading to a rise in overall consumption. This disparity is intensified by the varying degrees of loss aversion between the two groups, with high-income households feeling the pinch of profit reductions more acutely. Its amplification in recessions runs through the savers' lower effective curvature ($\sigma_r^H < \sigma_e^H$), not through any rise in loss aversion. Moreover, the effects of tax changes on low-income households are particularly severe during recessions, as the same regime shift in effective curvature makes them more sensitive to tax fluctuations.

NKPC The New Keynesian Phillips Curve is derived by:

$$\pi_t = E_t \beta \pi_{t+1} + \kappa_\zeta c_t + \kappa_\zeta \kappa_{g,\zeta} g_t - \kappa a_t, \quad \text{if } \zeta = [e, r], \tag{24}$$

where $\kappa = (\frac{1}{\Theta} - 1)(1 - \beta\Theta)(\phi + 1)$ is the standard Phillips-curve slope (Θ the Calvo probability of no price change), and the state-dependent slope κ_ζ and government-spending pass-through $\kappa_{g,\zeta}$ are the type-aggregated coefficients derived in the online Appendix B.5. Because loss aversion lowers both types' effective curvature in the loss regime, the population-weighted curvature entering the slope is smaller in recessions, so $\kappa_e > \kappa_r$.

During recessions, the Phillips Curve flattens as loss-averse preference increases willingness to work for both households, boosting substitutability between leisure and consumption. Our model, building on Santoro et al. (2014), shows that the NKPC's slope varies with household proportions and their degree of loss aversion, particularly under economic contractions when output becomes more sensitive to real interest rate changes. Furthermore, the parameter $\kappa_{g,\zeta}$ quantifies the impact of government spending on inflation, specifically reflecting the price increases stemming from labor market adjustments due to government spending, in line with the mechanism in Bilbiie (2025).

4.2 Monetary policy, Fiscal policy and inequality

Finally, we focus on the interaction between different policies and consumption inequality by first showing how the output affects the inequality via the general equilibrium channel from the output. We later provide an analytical way of showing how inequality shapes the demand market by visiting the IS curve and inequality.

Monetary policy To begin our analysis, we assume there is no government spending as before. Combining Equation (14) and goods market condition, $hc^H + (1-h)c^L = y$, after rearrangement, we can obtain the difference between the consumption:

$$c^H_{\zeta,t} - c^L_{\zeta,t} = \frac{1 - \chi^\zeta}{h} y_{\zeta,t}, \quad \text{if } \zeta = [e, r]. \tag{25}$$

As Bilbiie (2020) points out, inequality demonstrates counter-cyclical behavior, worsening during recessions. This relationship is captured by $\chi^e > 1$, indicating that increased output reduces inequality. Equation (25) further suggests this negative output-inequality relationship intensifies during recessions due to the higher consumption response among low-income individuals.

Our result coincides with the previous literature on monetary policy's impact on inequality. Specifically, the heterogeneous consumption responses to wage changes align with the prevailing belief that recessions hurt the poor the most, as argued by Guvenen et al. (2014), McKay and Wolf (2023). Consequently, if the government raises the interest rate, leading to a drop in output,

the poor are disproportionately affected. This transmission from interest rate to output during recessions intensifies the impact of contractionary monetary policy on increasing inequality.

To elucidate the impact of inequality on monetary policy within our IS curve framework, consider the interest rate coefficient $\psi_{c,\zeta} = \frac{h}{1-(1-h)\chi^{\zeta}}$, along with the consumption differential between high and low-income groups, defined as $c^H_{\zeta,t} - c^L_{\zeta,t} = \frac{1-\chi^{\zeta}}{h} y_{\zeta,t}$. In scenarios where low-income households hold fewer dividend claims and constitute a larger share of the population, they significantly drive aggregate demand due to their higher propensity to consume rather than save or invest.

As inequality diminishes, a similar increase in aggregate demand occurs. This increase is driven by a greater share of consumption shifting to lower-income groups, who spend a higher fraction of their income. Consequently, when monetary policy actions, such as interest rate cuts, reduce inequality, they also amplify the policy's effectiveness by bolstering aggregate demand.

This dynamic underscores the interplay between monetary policy and economic inequality, significantly influencing policy efficacy throughout various stages of the economic cycle.

Fiscal policy Let us now turn to examine how fiscal policy influences inequality. After accounting for consumption, $c = y - g$ represents total consumption, we derive:

$$c^H_{\zeta,t} - c^L_{\zeta,t} = \frac{1-\chi^{\zeta}}{h} c_{\zeta,t} - \psi_{g,\zeta} g_t + \psi_{\tau,\zeta} t^L_t, \quad \text{if } \zeta = [e, r]. \tag{26}$$

At first glance, we see that inequality remains counter-cyclical as before, adding government spending does not change this property. Indeed, government spending introduces two additional components influencing inequality: the government expenditure itself and the relative tax burden on low-income households compared to the high-income. These components align with the consumption response of low-income households.

The coefficient $\psi_{g,\zeta}$ reflects the differing savings responses when facing the associated cost of government spending between two income groups. High-income groups anticipate tax increases or future debt repayment more, adjusting their savings accordingly due to their larger share of firm profits. This anticipatory saving is adjusted for the high-income population share and multiplied by the consumption response of low-income households.

For the tax burden, if low-income individuals shoulder a larger share of government spending through taxes, inequality will widen, determined by labor elasticity. During a recession, this is captured by $\psi_{\tau,\zeta}$. During expansions, taxes affect low-income individuals' budget constraints without a general equilibrium effect, but during recessions, high-income loss aversion causes

increased consumption inequality as low-income tax burdens affect high-income consumption relatively more, observed in the ratio $\frac{\chi^r}{\chi^e}$.

The correspondence between the IS curve coefficient for fiscal policy and consumption inequality, as shown in Equation (26), links fiscal policy's effects on overall consumption and inequality. Government spending, represented by $\psi_{g,\zeta}$, affects consumption levels and modulates inequality, influenced by the proportion of low-income consumers and the aggregate policy impact $(\psi_{c,\zeta})$.

An increase in government spending prompts both income groups to engage in anticipatory savings, expecting future tax hikes. High-income households, facing greater idiosyncratic risks due to their larger share of dividend claims and higher loss aversion, tend to save more. In contrast, low-income households consume a greater portion of their income, driven by a smaller inclination towards such anticipatory saving. The multiplier effect of this behavior, encapsulated in $\psi_{g,\zeta}$, is compounded by their consumption response to government spending.

This implies that the stimulative effect of fiscal policy on aggregate demand and its capacity to reduce inequality is magnified in economic downturns. This result aligns with empirical findings highlighting the enhanced effectiveness of fiscal interventions during recessive periods (e.g., Auerbach and Gorodnichenko (2012a) and Ghassibe and Zanetti (2022)). The model illuminates the nuanced role of fiscal policy in influencing economic output and inequality under varying economic conditions, emphasizing the importance of context-specific policy design and implementation.

4.3 Asymmetric Impact on Optimal Policy

The preceding analysis has shown that loss aversion generates asymmetric consumption responses across business cycles: low-income households' consumption becomes more sensitive to output during recessions, amplifying policy effects on both inequality and general equilibrium. This asymmetry has critical welfare implications. A welfare-maximizing central bank should tolerate more inflation during recessions than expansions because the flatter Phillips curve and amplified inequality response make disinflation through output contraction particularly costly, with losses falling disproportionately on low-income households. Hence, a "state-blind" central bank that ignores this asymmetry makes systematic errors during recessions will fail to lower the interest rate enough, generating substantial welfare losses that exacerbate inequality and feed back through general equilibrium effects, making such policy suboptimal.

To make the argument precise we follow Bilbiie (2025) and write the welfare criterion, ab-

sent government spending, as a state-dependent loss over inflation, output, and consumption inequality; substituting the equilibrium relation $\gamma_t = \frac{1-\chi^\zeta}{h} y_t$ reduces it to $\mathcal{L}^\zeta = \frac{1}{2} E_0 \sum_t \beta^t \left[\pi_t^2 + \Lambda^\zeta y_t^2\right]$, with a composite output weight Λ^ζ that is larger in recessions because the flatter Phillips curve and the amplified inequality response both raise the cost of stabilizing inflation through output (the full derivation, including the welfare weights and the state-dependent targeting and interest-rate rules, is in the online Appendix E). The optimal discretionary policy is therefore *less* aggressive against inflation in recessions, and the state-dependent interest-rate response to expected inflation satisfies $\phi^r < \phi^e$.

But why should central bankers care? A policymaker who ignores the state-dependence and applies the expansion coefficient ϕ^e during recessions over-tightens. The policy coefficient gap can be decomposed as:

$$\Delta\phi \equiv \phi^e - \phi^r = \underbrace{\sigma_e^H \kappa^e \left[\frac{1}{\psi_{c,e}\Lambda^e} - \frac{1}{\psi_{c,r}\Lambda^r}\right]}_{\text{GE effect}} + \underbrace{\frac{\sigma_e^H(\kappa^e - \kappa^r)}{\psi_{c,r}\Lambda^r}}_{\text{Phillips curve effect}} + \underbrace{\frac{(\sigma_e^H - \sigma_r^H)\kappa^r}{\psi_{c,r}\Lambda^r}}_{\text{IES effect}}. \tag{27}$$

The first term captures the general equilibrium effect: the GE multiplier is larger in recessions ($\psi_{c,r} > \psi_{c,e}$) and policymakers must place greater weight on output ($\Lambda^r > \Lambda^e$). The second term reflects the flattening of the Phillips curve in recessions, which makes disinflation more costly in terms of output sacrifice. The third term arises from loss aversion: in the loss regime the savers' effective curvature is lower ($\sigma_r^H < \sigma_e^H$), so a given interest rate move shifts demand by more. All three terms are positive, confirming $\Delta\phi > 0$.

Using ϕ^e instead of ϕ^r creates excessive policy tightening during recessions, generating welfare losses through both higher inflation volatility and amplified output contractions. These losses are particularly severe during cost-push shocks such as those experienced during the COVID-19 pandemic, when ignoring loss aversion and state-dependent inequality dynamics leads to systematic policy mistakes that disproportionately harm low-income households. Because the linear-quadratic approximation is taken regime by regime around a deterministic reference point, we report the direction and composition of the welfare losses rather than a consumption-equivalent magnitude.

So far, we have only focused on analytical insight. But does this model really capture the empirical evidence we have? To address this question, we perform quantitative exercises in the next section.

5 Quantitative Analysis

We perform our quantitative analysis with the full model, a TANK model with type switching (Bilbiie, 2020), allowing the probability of switching types for each period.[20] To conduct our quantitative analysis with precision, we calibrate the model parameters, targeting key macroeconomic data that we will elaborate on subsequently. This calibration process aims to ensure our model accurately reflects real-world economic conditions. Upon setting these parameters, we generate simulated data over 5000 impulse response periods. These simulations are then treated as if they were actual real-world data. We apply the same local projection method in Equation 1 for a direct comparison to the empirical results. With two types the simulated "Gini" of any variable is proportional to the population-weighted type gap $h(1-h)(x_t^H - x_t^L)$ in log-deviation units; the proportionality constant is common to both states, so it does not affect the recession-expansion comparison, and the alternative-shock comparison of the online Appendix C.1 addresses the units formally.

In the simulation a quarter is a loss (recession) regime when aggregate consumption is below its reference ($\delta_t < 0$), held fixed within each impulse-response window. Because the technology process is persistent, loss regimes are more frequent than the NBER contractions that date the empirical states (roughly half of simulated quarters versus about 16 percent).

5.1 Calibration

We present our calibration values in Table 1 and begin with standard parameters used in the New Keynesian model. We set the discount rate, $\beta = 0.99$, consistent with a 4% annual interest rate as commonly assumed. Following precedents in macroeconomic literature, we set the Frisch elasticity, $\phi = 1$ in line with Bilbiie (2020). To represent demographic distributions, we calibrate the proportion of the hand-to-mouth at 25%, aligning with estimation in Havranek and Sokolova (2020). We set the dividend share of hand-to-mouth households, $\mu^d = 0.12$, implying that high-income households hold about 88% of dividend shares, resonating with the findings of Kuhn et al. (2020). We calibrate the adherence probability to the same agent type, $s = 0.96$, ensuring an 85% annual probability for agents to maintain their type consistent with McKay et al. (2016); the low-type persistence s_L adjusts to hold $1 - h = (1-s)/(2-s-s_L) = 0.25$ (so $s_L = 0.88$), the

[20] The derivation and results of the full model to perform simulation are shown in the online Appendix B. In the online Appendix C, we also simulate endogenous reference points based on Santoro et al. (2014). Additionally, we follow Santoro et al. (2014) to ensure policy shock is not too large to change the economic state.

same convention applying when s varies by state in the online Appendix D.1.

Next, we set the loss aversion coefficient for high-income savers, λ^H, at 2.75, at the upper end of the individual-level estimates of Tversky and Kahneman (1992) and Gächter et al. (2022), consistent with the saver type being drawn from the more loss-averse part of the population. For hand-to-mouth households, we target the consumption response around gains and losses reported by Fuster et al. (2021). We calibrate the loss aversion coefficient for hand-to-mouth households, $\lambda^L = 1.25$, to target the consumption response of the output, $\chi^r = 1.6$ estimated by Fuster et al. (2021) on the consumption respond to output during losses of \$500.[21] Finally, we adopt values from Santoro et al. (2014), setting the sensitivity of loss aversion, $\theta = 1$, and the preference for loss-aversion, $\alpha = 0.5$. Because λ^L is chosen to reproduce the Fuster et al. (2021) loss response, the model's ability to match Fuster-style responses is a property of the calibration, not an independent validation; the validation content lies in the untargeted local-projection asymmetries.

On the supply side, the price markup is set at 1.2, with the probability of price changes being 85%, in line with standard New Keynesian literature for Calvo pricing (see e.g., Calvo (1983)). This setting yields a standard Phillips Curve slope, $\kappa = 0.06$, which we adjust for expansion ($\kappa_e = 0.07$) and recession ($\kappa_r = 0.05$) periods.[22] The persistence of the technology shock, ρ_a, is calibrated at 0.9, with a variance of $\sigma_a^2 = 0.06$, matching the volatility observed in five-year earnings distributions as reported by McKay et al. (2016).

For baseline policy parameters, we choose the following: The variance of monetary policy shock $\sigma_m^2 = 0.02$, aligning with Santoro et al. (2014) with persistence, ρ_m, set at 0.8 as used in Bilbiie (2020),[23] and the interest rate response to inflation, $\phi_\pi = 1.5$, is derived from the standard Taylor rule. The fiscal policy shock resistance, $\rho_g = 0.9$, with a variance of $\sigma_g^2 = 0.2$, consistent with estimations in Blanchard and Perotti (2002) and used in Galí et al. (2007). Finally, the fiscal policy role $\phi_g = 0.06$ is calibrated to match the average 6% net government spending to GDP ratio, and we assume a uniform tax rate. In the zero-liquidity equilibrium ($B^G = 0$) the debt-feedback rule reduces to budget balance ($g_t = t_t$), so ϕ_g matters only through the steady-state spending share.

[21] Hence, the overall loss-aversion in the society is 2.38 which is similar of the loss-aversion value 2.25 applied in Santoro et al. (2014) under the representative agent model.

[22] Precisely $\kappa = (1/\Theta - 1)(1 - \beta\Theta)(1 + \phi) = 0.056$, so $\kappa_e = 0.070$ and $\kappa_r = 0.049$; the online Appendix B states the aggregation behind κ_r.

[23] Santoro et al. (2014) show this value will prevent the monetary policy shock to switch economic state.

Table 1: Calibration of Parameters

Parameter	Description	Value
λ^H	High-Income Loss aversion	2.75
λ^L	Low-Income Loss aversion	1.25
θ	Sensitivity to Loss Aversion	1
α	Preference to Loss Aversion	0.5
β	Discount Rate	0.99
ϕ	Inverse of Frisch elasticity	1
κ	New Keynesian Philips Curve Slope	0.06
μ_d	Share of Dividend of the low-income	0.12
ρ_m	Autocorrelation of monetary shock	0.8
ρ_a	Autocorrelation of technology shock	0.9
ρ_g	Autocorrelation of fiscal shock	0.9
σ_m^2	Variance of monetary shock	0.02
σ_a^2	Variance of technology shock	0.06
σ_g^2	Variance of fiscal shock	0.2
ϕ_g	Government spending ratio	0.06

Note: This table provides the calibration value of parameters used in the model simulation.

5.2 Results

We present our main simulation result in Figure 6 for policy impact on consumption inequality across the business cycle. In these illustrations, the empirical results derived from prior analyses are denoted by the red dotted line, and the outcomes from our model simulations are represented by the blue line. A critical observation from our analysis is the pronounced impact of policy during recessions in contrast to expansions.

Table 2 presents a detailed comparison of model-generated and empirical cumulative impulse responses to contractionary monetary and expansionary fiscal policy shocks over horizons from one to five years (4, 8, 12, and 20 quarters). Columns 3 and 6 report the model's accuracy in matching the empirical outcomes, as the ratio of empirical and simulation results for monetary and fiscal policy respectively.

In the AR(2) setting for monetary policy, initial alignment with empirical observations is strong at 86% in the first year, increasing to 97% in the second, but then drops to 51% by the

Figure 6: Comparing Empirical and Simulated Results for Consumption Inequality



Note: The figure illustrates the differential impacts of policy measures during recessions and expansions, quantified as $\beta^r - \beta^e$, in response to a 1% contractionary monetary policy shock (left panel) and a 1% expansionary fiscal policy shock (right panel) on consumption inequality. This analysis employs the local projection method with AR(2) and MA(20) specified in Eq. (1) and simulates the model over 5,000 impulse response periods, represented by the blue line. The empirical results of Section 2.2 (Romer and Romer (2004) monetary and Blanchard and Perotti (2002) fiscal shocks, 1980-2008, smooth-transition state) are shown by the red dotted line with confidence intervals (green vertical lines).

third year and 42% by the fifth year, indicating a decrease in predictive accuracy over longer periods. For fiscal policy under AR(2), the model starts with a robust 74% correspondence with empirical data but drops to 43% in the second year, stabilizes at 31% in the third, and rebounds to 50% by the fifth year. On balance the model accounts for 50 to 80 percent of the asymmetric consumption-inequality responses, and it also delivers the paper's central fact: a state-dependent response of consumption inequality with a near-flat response of relative earnings. In the online Appendix C.1 (Figure C.2) we present our findings on earnings inequality[24], which we define in our model as the sum of the wage and dividend return—closer to total income than to labor earnings; the effects of policy on earnings inequality are relatively consistent across different phases of the business cycle, in line with our empirical findings, and empiri-

[24] Additionally, in the online Appendix C we also provide results for output and inflation.

Table 2: Asymmetric Policies Impact on Inequality for Empirical and Theoretical IRFs

	Monetary Policy ($\beta^r - \beta^e$)			Fiscal Policy ($\beta^r - \beta^e$)		
Time (quarters)	Model (in %)	Data (%)	Ratio (%)	Model (%)	Data (%)	Ratio (%)
AR2						
4	1.16	1.35	85.90	-1.78	-1.42	74.08
8	1.42	1.47	96.73	-4.02	-1.65	42.81
12	1.52	2.99	50.88	-6.53	-2.82	31.32
20	1.49	3.58	41.59	-8.71	-5.80	49.77
AR1						
4	1.16	0.47	46.53	-1.77	-1.61	90.27
8	1.43	0.63	25.70	-4.00	-1.43	79.48
12	1.52	1.16	69.22	-6.49	-2.28	84.80
20	1.50	1.31	85.77	-8.65	-5.04	28.17

Note: This table displays the cumulative impulse responses to a contractionary monetary policy shock (left panel) and an expansionary fiscal policy shock (right panel) across four time horizons: 1 year, 2 years, 3 years, and 5 years. The "Model" column details responses as percentages derived from 5000 iterations using the local projection method in Eq. (1), while the "Data" column shows responses based on empirical findings. The ratios reported indicate the percentage deviation between model outcomes and empirical results, calculated as $1 - \frac{|Model - Data|}{Data}$. Analysis is split into AR(2) specification (upper panel) and AR(1) specification (lower panel).

cally both labor earnings and total before-tax income show no robust state dependence, so the comparison holds under either definition.

Significant discrepancies between the simulated and empirical data appear after 10 periods following a monetary policy shock. These differences might be due to the use of an AR(1) process in our shock process, unlike the AR(2) process used in empirical projections. In the online Appendix C, we depict the result of AR(1) specification for local projection and this shows more accurately mirrors the empirical data concerning consumption inequality subsequent to a monetary policy shock after 10 periods.

Discussion In this section, we discuss why conventional models may struggle to generate the patterns observed in our empirical results, particularly the asymmetric redistribution of policy effects on consumption across economic states while earnings inequality remains relatively stable.

The traditional Two-Agent New Keynesian (TANK) framework (see, e.g., Campbell and

Mankiw (1991) and Galí et al. (2007)) typically attributes policy effects on inequality to differences in income sources and budget constraints. In this setup, variations in consumption across economic states are primarily driven by differences in income, making it difficult to explain why consumption inequality worsens during recessions while earnings inequality remains relatively symmetric.

However, recent studies in the Heterogeneous-Agent New Keynesian (HANK) literature challenge the sufficiency of income differences in explaining inequality dynamics, instead emphasizing the roles of consumption and wealth channels (see, e.g., Kaplan et al. (2014), Kaplan et al. (2018), Kaplan and Violante (2022), Ahn et al. (2018), and Bilbiie (2020)).[25] While our model is based on loss aversion, an alternative explanation from the HANK literature is countercyclical income risk. For instance, Guvenen et al. (2014) find that high-income households face a higher probability of income loss during recessions, which strengthens their precautionary motives in incomplete markets. Recent heterogeneous models such as Bilbiie (2025) and Auclert (2019) incorporate this feature into the HANK framework to generate distributional effects through precautionary savings.

In the online Appendix D, we examine this alternative explanation in detail and show that, for reasonable calibrations of countercyclical income risk, the HANK model with idiosyncratic uninsured income risk alone cannot quantitatively account for the observed patterns of consumption and earnings inequality over the business cycle. This discrepancy arises because loss aversion affects both consumption-labour choices and intertemporal decisions, whereas income risk primarily influences precautionary savings.

6 Conclusion

We have documented that contractionary monetary and fiscal policy raise consumption inequality more strongly in recessions, with no robust counterpart in earnings inequality. Our model, based on Bilbiie (2020), analytically explains this pattern through heterogeneity in loss aversion. When consumption falls below its reference point in a recession, households become more responsive to policy, as in Santoro et al. (2014); because high-income households are more loss-averse, they are more reluctant to let their consumption fall and smooth it by working more, so

[25] For example, Kaplan et al. (2018) show that households' responses to output changes vary based on their liquidity levels, with those holding fewer liquid assets reacting more strongly. Additionally, Bilbiie (2020) highlights the role of idiosyncratic uninsured income risk in generating heterogeneous consumption responses among income groups.

the burden of the adjustment falls on low-income households, who are more reliant on wages and whose consumption is more sensitive to current income. These mechanisms explain why both monetary and fiscal policy have stronger distributional effects during recessions, and why the pattern appears in consumption but not earnings inequality.

Our model does not include the financial market, although Kaplan et al. (2014) and Ahn et al. (2018) highlight the role of market heterogeneity in influencing inequality, and Li and Yang (2013) and Ang et al. (2005) indicate that loss aversion significantly affects financial markets. Future research could enrich this model by incorporating a detailed household finance structure to analyze how loss aversion shapes inequality over the business cycle.

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Online Appendix

Appendix for online publication only

A Additional Empirical Results

In this section, we present additional empirical results that are not included in the main text due to space constraints. The order follows a single logic: we first report the baseline results in full, then vary one ingredient at a time. Throughout, inference tables report the cumulated one- to five-year horizons of Table 2 ($h = 4, 8, 12, 20$ quarters), which bracket the two- to three-year window in which the real effects of monetary policy shocks peak (Romer and Romer (2004)), and follow the practice of Ramey and Zubairy (2018) of reporting cumulative estimates at fixed horizons in tables with the full paths in figures; Newey-West standard errors are in parentheses with significance stars at the 10, 5, and 1 percent levels (truncation lag $h + 1$, since cumulated local-projection residuals are serially correlated by construction; the baseline smooth-transition design of Figures 1–3 and Table **??** instead reports Driscoll-Kraay inference, which is robust to serial correlation and to the correlation across horizons in the pooled regression, as in Section 2). Appendix A.1 details the construction of the monetary and fiscal instruments and their aggregate (state-independent) macro responses, and Appendix A.2 reports the baseline inequality results across measures; both extend the analysis of Coibion et al. (2017) to additional instruments and to fiscal shocks. Appendix A.3 presents the state-dependent results across inequality measures and both state definitions and addresses the associated econometric concerns. We then vary the identification and sample: Appendix A.4 re-estimates the baseline on NBER dating and alternative shock series, Appendix A.5 presents alternative monetary series across identification families, and Appendix A.6 collects the per-series point estimates behind the figures. A note on confidence-band conventions: the figures of Appendices A.1–A.3 are retained from earlier vintages of this project and display the band stated in each figure note (one-standard-error bands in several legacy figures), whereas every main-text figure and every figure in Appendices A.4–A.6 reports 90 percent bands; no significance claim in the text rests on a legacy figure's bands.

A.1 Policy Shock and Macro Variables

Estimation of Monetary Policy Shock Narrative identification is our preferred approach, for the reasons emphasized by Aruoba and Drechsel (2024): it controls directly for the policymaker's information set—through the Greenbook forecasts in Romer and Romer (2004) and through the

FOMC's own documents in Aruoba and Drechsel (2024)—so the identified innovation is purged of the central bank's systematic response to, and private information about, the economy, the concern that Jarociński and Karadi (2020) and Bauer and Swanson (2023) raise for high-frequency surprises; and its samples reach back to 1980 and earlier, covering more recessions, which matters for a state-dependent design in which recessions are scarce. The main result of the estimation of monetary policy is adopted from Romer and Romer (2004) for the identification of monetary policy innovations, separating them from anticipatory effects linked to economic conditions. Their methodology involves the creation of a historical record detailing shifts in the target Federal Funds rate (FFR) during each Federal Open Market Committee (FOMC) meeting from 1969 to 1996. Additionally, they generate a measure of monetary policy shocks based on the deviations in policy changes at each meeting that do not align with the information available to the Fed at the time. Specifically, their estimation encompasses

$$\begin{aligned}\Delta ff_m =&\alpha + \beta ffb_m + \sum_{i=-1}^{2} \gamma_i F_m \Delta y_{m,i} + \sum_{i=-1}^{2} \lambda_i \left(F_m \Delta y_{m,i} - F_{m-1}\Delta y_{m,i}\right) + \sum_{i=-1}^{2} \varphi_i F_m \pi_{m,i} \\ &+ \sum_{i=-1}^{2} \theta_i \left(F_m \pi_{m,i} - F_{m-1}\pi_{m,i}\right) + \mu_i F_m ue_0 + \varepsilon_m,\end{aligned} \tag{A.1}$$

where m denotes the FOMC meeting, ffb_m is the target FFR going into the FOMC meeting, $F_m \Delta y_{m,i}$ is the Greenbook forecast from meeting m of real output growth in quarters around meeting m, $F_m \pi_{m,i}$ are Greenbook forecasts of GDP deflator inflation, and $F_m ue_0$ are Greenbook forecasts of the current quarter's average unemployment rate. The monetary policy shock is the estimated residual $\hat{\varepsilon}_m$. (The symbols in this regression follow the notation of Romer and Romer (2004) and are local to this equation; they are unrelated to the model parameters of the main text.)

To further validate our result, we also use the monetary policy shock from Aruoba and Drechsel (2024). Aruoba and Drechsel (2024) shock data period is closer to Romer and Romer (2004) and conceptually in line with Romer and Romer (2004) using the narrative approach to identify monetary policy shock. The difference is that Aruoba and Drechsel (2024) adopt textual analysis instead of the Greenbook forecasts used in Romer and Romer (2004), obtaining macroeconomic information from Fed staff documents that is not covered in the Greenbook. The narrative series end in 2008; to extend the sample and the identification set, the online Appendix A.5 re-estimates the results with the high-frequency families (Jarociński and Karadi (2020), Gertler and Karadi (2015), Swanson (2021), and Bauer and Swanson (2023)), and the results are qualitatively unchanged.

Estimation of Fiscal Policy Shock The standard local projection method of Jordà (2005) is based on the equation:

$$x_{t+h} = \alpha_h + \psi_h(L) z_{t-1} + \beta_h \, \text{shock}_t + \varepsilon_{t+h} \quad \text{for } h = 0, 1, 2, \ldots, \tag{A.2}$$

where $\psi_h(L)$ is a polynomial in the lag operator L (not to be confused with the low-income type index).

When applying the Blanchard-Perotti identification approach, the shock is straightforwardly determined by the current government spending since the set of controls, denoted as z, encompasses lagged measurements of both GDP and government spending. Consequently, this aligns with the Blanchard-Perotti structural VAR identification methodology. The coefficient β_h signifies the reaction of variable x at time $t+h$ to the shock that occurs at time t. Thus, the process of constructing impulse responses involves generating a series of β_h estimates through individual regressions for each time horizon.

The availability of the shock series varies: Aruoba and Drechsel (2024) is available from 1982/10 to 2008/12, and Romer and Romer (2004) from 1969/01 to 2008/12 (we use observations from 1980/01, as in the main text). The advantage of the differing time periods is that they allow us to check whether the results are driven by particular severe recessions or are general across business cycles. We aggregate these monthly shocks to a quarterly level by summing the three monthly values.

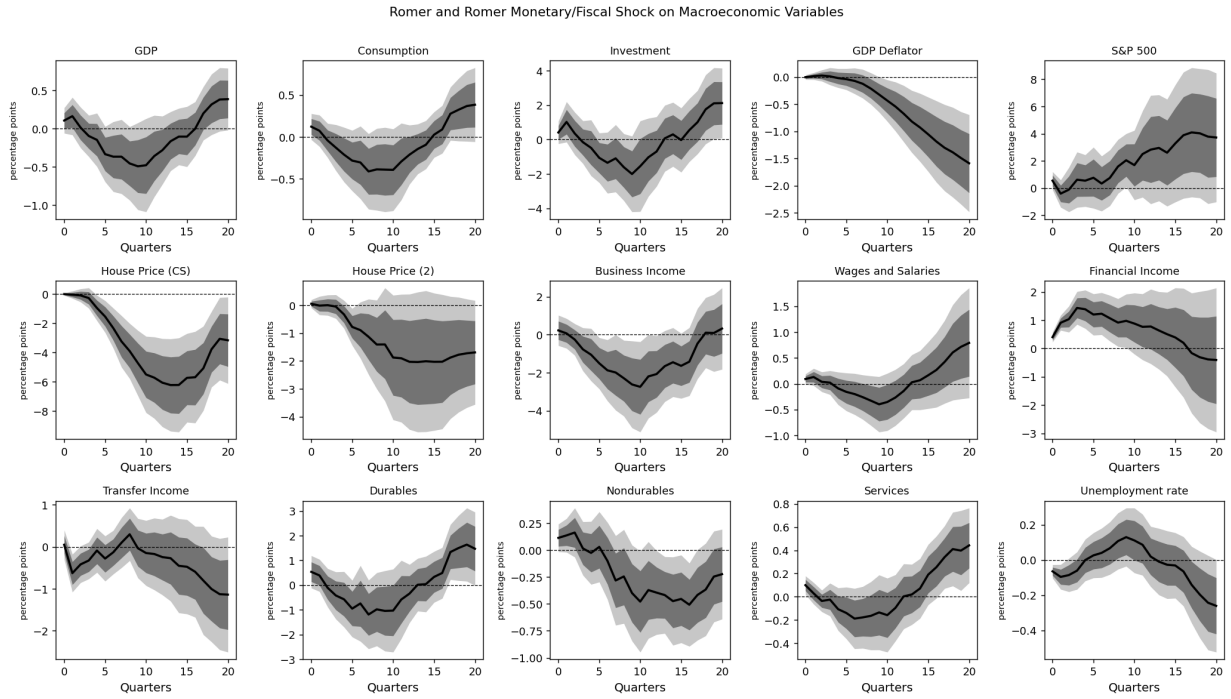
We use the local projection method to estimate the impact of monetary policy shocks on various macroeconomic variables. The regression model is specified as follows:

$$x_{t+h} - x_{t+h-1} = c^{(h)} + \sum_{j=1}^{J} \alpha_j^{(h)} \left(x_{t-j} - x_{t-j-1} \right) + \sum_{i=0}^{I} \beta_i^{(h)} \eta_{t-i} + \varepsilon_{t+h}, h = 0, \ldots, H. \tag{A.3}$$

In this equation, we estimate the effect of the policy shock η on changes in the outcome variable x over different horizons h. We control for J lags of the dependent variable and I lags of the policy shock. Following Coibion et al. (2017), we set $J = 2$ and $I = 20$, covering 20 horizons.

Figure A.1 replicates the result of Coibion et al. (2017), with one-standard-error confidence intervals, as we use the same data period and methodology. In Figure A.2 we present the result using the Aruoba and Drechsel (2024) monetary policy shock, and the responses are very close to those under Romer and Romer (2004). For fiscal policy, we use the Blanchard and Perotti (2002)

Figure A.1: Romer and Romer Monetary Shock on Macroeconomic Variables



Note: This figure shows the impact of 1 standard deviation contractionary Romer and Romer (2004) monetary policy shock on macro variable measured by the change of the log as defined in Coibion et al. (2017). Housing price is measured by the Case-Shiller index(i.e. CS) and city housing price data as obtained from Coibion et al. (2017). Using AR2 and MA20 local projection method. Data period from 1980/01 to 2008/12.

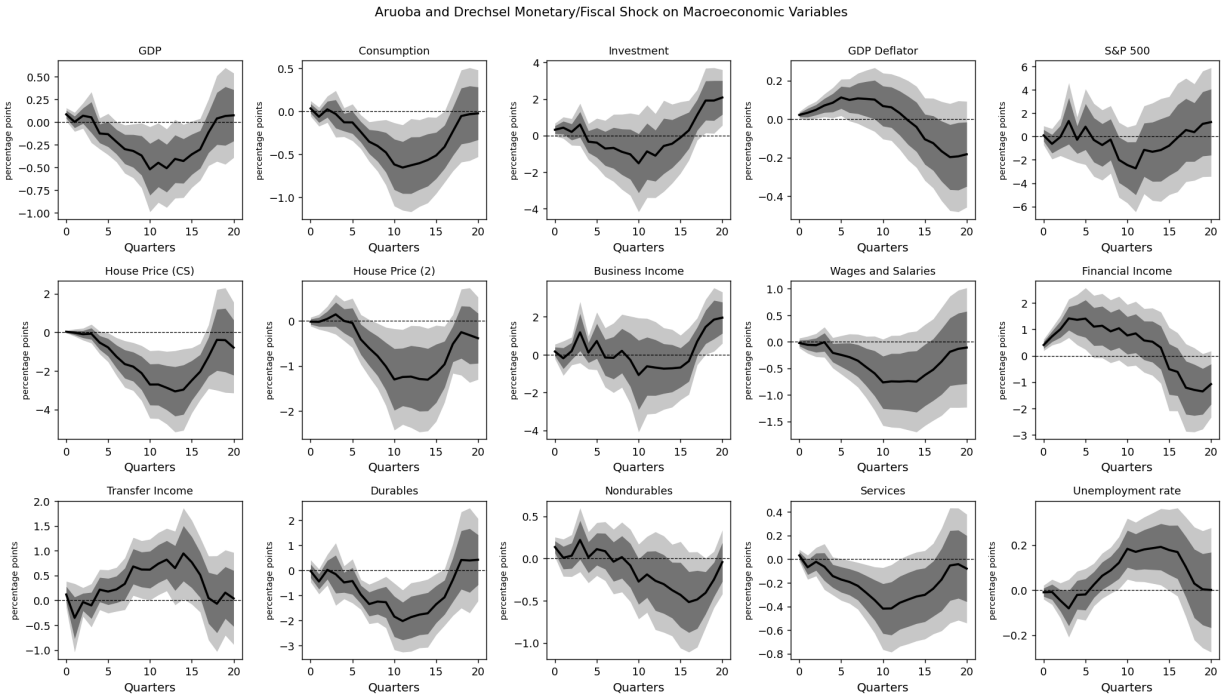
shock, following the approach in Ramey and Zubairy (2018) and the result is shown in Figure A.3.

A.2 Policy Shock and Inequality

We compare the effects of different monetary and fiscal policy shocks on inequality using the local projection method, based on the CEX inequality series described in the main text. As noted by Coibion et al. (2017), the survey's monthly data are aggregated quarterly, creating cohorts that may induce cross-sectional serial correlations. To address this, we use Driscoll-Kraay standard errors, which adjust for these correlations under the assumption that policy shocks do not immediately impact inequality.

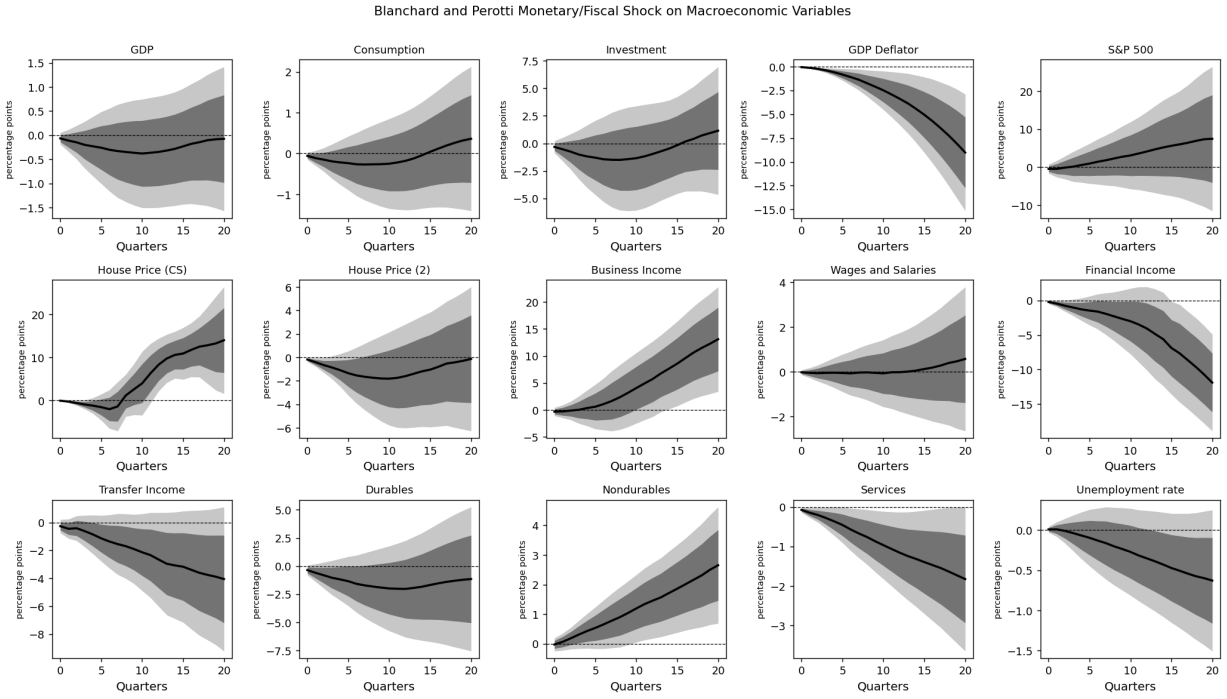
Alternatively, if Driscoll-Kraay standard errors are not used, we can treat each surveyed household cohort separately, averaging their results with Newey-West standard errors. However, this approach limits the sample size to one-third of the total, leading to noisier inequality measures compared to those calculated for the entire sample (see further discussion in Coibion

Figure A.2: Aruoba and Drechsel Monetary Shock on Macroeconomic Variables



Note: This figure shows the impact of 1 standard deviation contractionary Aruoba and Drechsel (2024) monetary policy shock on macro variable measured by the change of the log as defined in Coibion et al. (2017). Housing price is measured by the Case-Shiller index(i.e. CS) and city housing price data as obtained from Coibion et al. (2017). Using AR2 and MA20 local projection method. Data period from 1982/10 to 2008/12.

Figure A.3: Blanchard and Perotti Fiscal Shock on Macroeconomic Variables



Note: This figure shows the impact of 1 standard deviation expansionary Blanchard and Perotti (2002) fiscal policy shock estimated in Ramey and Zubairy (2018) on macro variable measured by the change of the log as defined in Coibion et al. (2017). Housing price is measured by the Case-Shiller index(i.e. CS) and city housing price data as obtained from Coibion et al. (2017). Using AR2 and MA20 local projection method. Data period from 1980/01 to 2008/12.

et al. (2017) Appendix A).

The results for monetary policy shocks are shown in Figures A.4, A.5, and A.6 for the Gini coefficient, 90-10 ratio, and standard deviation, respectively. We analyze the impact using shocks from Romer and Romer (2004) and Aruoba and Drechsel (2024).

The results indicate that income and earnings inequality show little variation across different shock measurements. However, for consumption and expenditure, the impact of the Romer and Romer (2004) shock appears larger, especially for expenditure inequality. This might be due to the central-bank information channel, which pushes consumption inequality higher. Overall, using alternative monetary policy shocks does not change the main results of Coibion et al. (2017) as we can see from the result.

The results of the fiscal policy impact are shown in Figure A.7, where we report the standard deviation, Gini coefficient, and 90-10 ratio as measures of inequality. Similar to monetary policy, the impact on income and earnings inequality is relatively muted. However, fiscal policy shocks tend to reduce consumption and expenditure inequality, except for the standard deviation of consumption, which shows less consistent results.

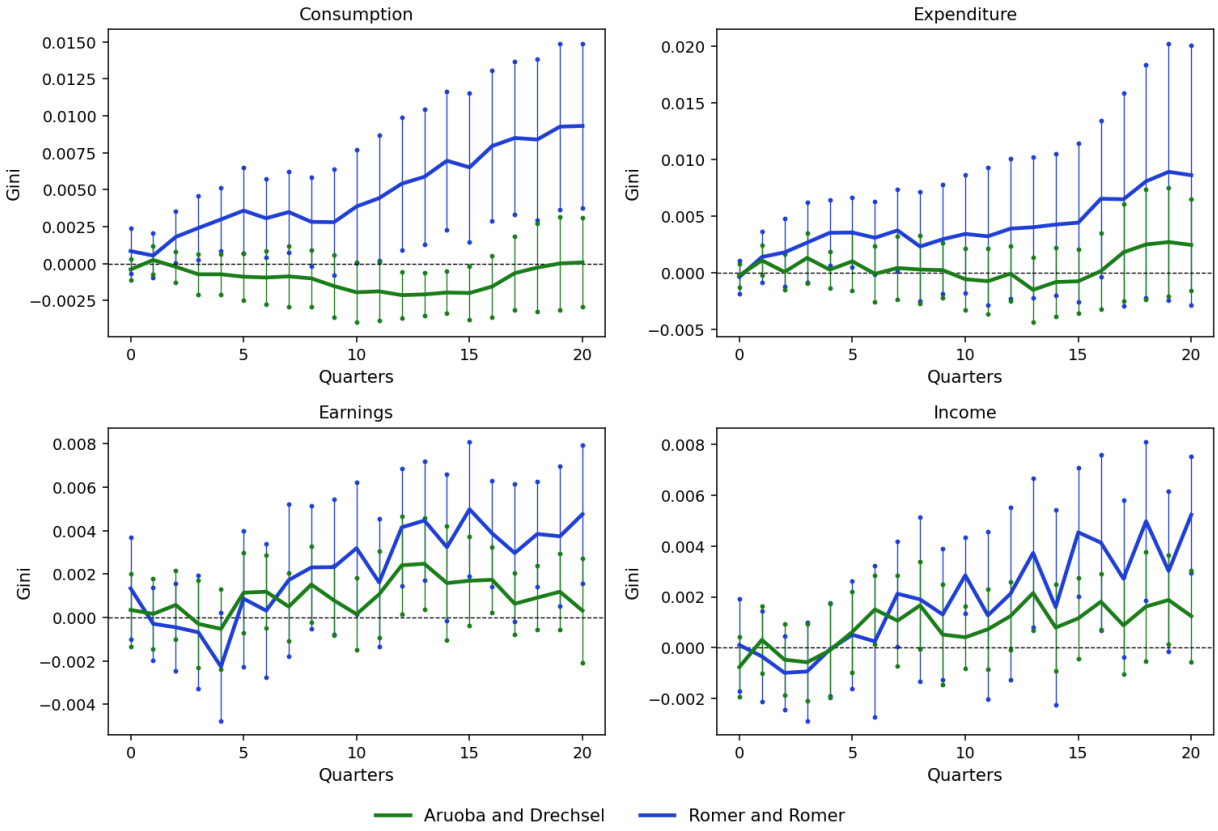
In the next section, we will further analyze the impact of policy on inequality by dividing it across different phases of the business cycle.

A.3 State-Dependence Policy Shock on Inequality

This appendix collects the complete set of state-dependent results under the main-text configuration (the Romer and Romer (2004) and Aruoba and Drechsel (2024) monetary shocks and the Blanchard and Perotti (2002) fiscal shock on 1980-2008, the smooth-transition recession indicator, and Driscoll-Kraay inference) for all four inequality measures (Gini, 90-10, and cross-sectional standard deviation of consumption, expenditure, earnings, and total income). In implementation, the main specification of Eq. (1) is estimated with the cumulated change in the inequality measure as the dependent variable:

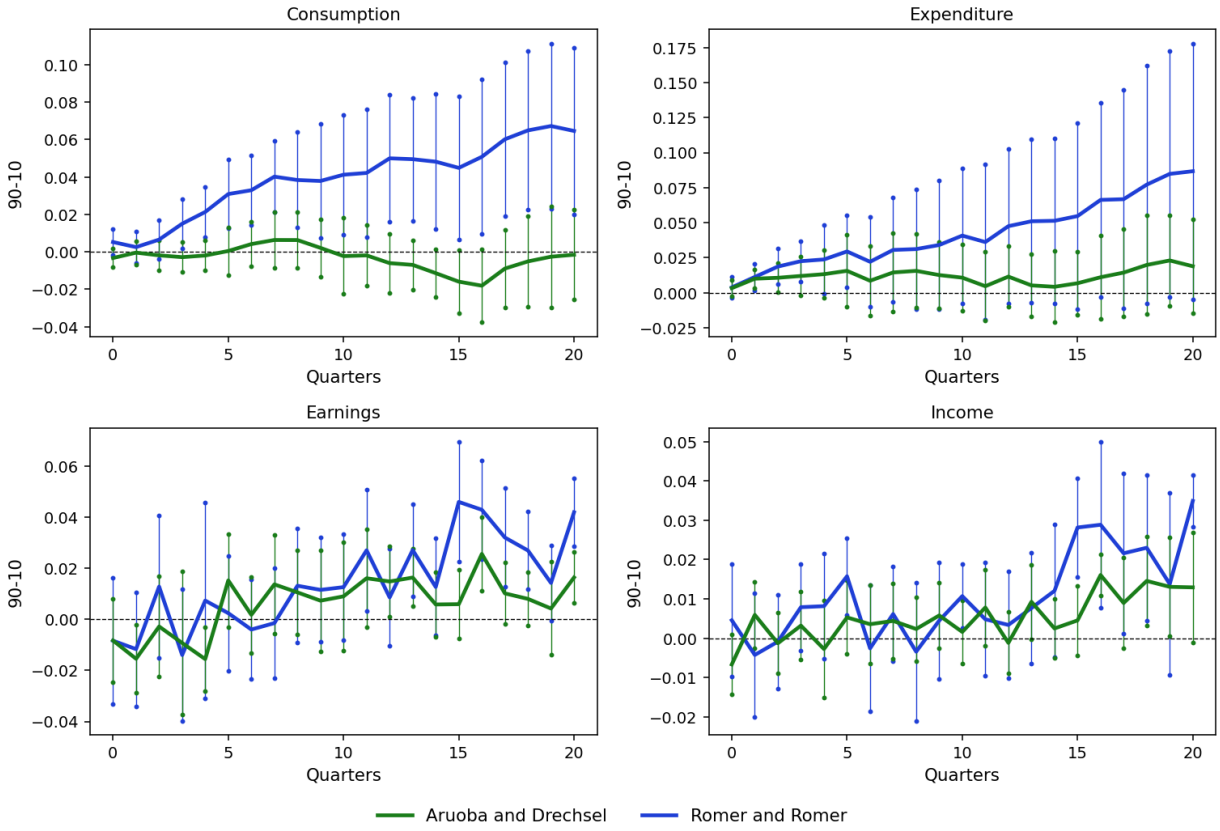
$$\begin{aligned} x_{t+h} - x_{t-1} = c^{(h)} &+ F(z_t)\left[\sum_{i=0}^{I}(\beta_i^r)^{(h)}\eta_{t-i} + \sum_{j=1}^{J}(\alpha_j^r)^{(h)}x_{t-j}\right] \\ &+ (1-F(z_t))\left[\sum_{i=0}^{I}(\beta_i^e)^{(h)}\eta_{t-i} + \sum_{j=1}^{J}(\alpha_j^e)^{(h)}x_{t-j}\right] + \varepsilon_{t+h}, h = 0, \\ &\ldots, H. \end{aligned} \tag{A.4}$$

Figure A.4: Different Monetary Policy Shocks on Inequality: Gini



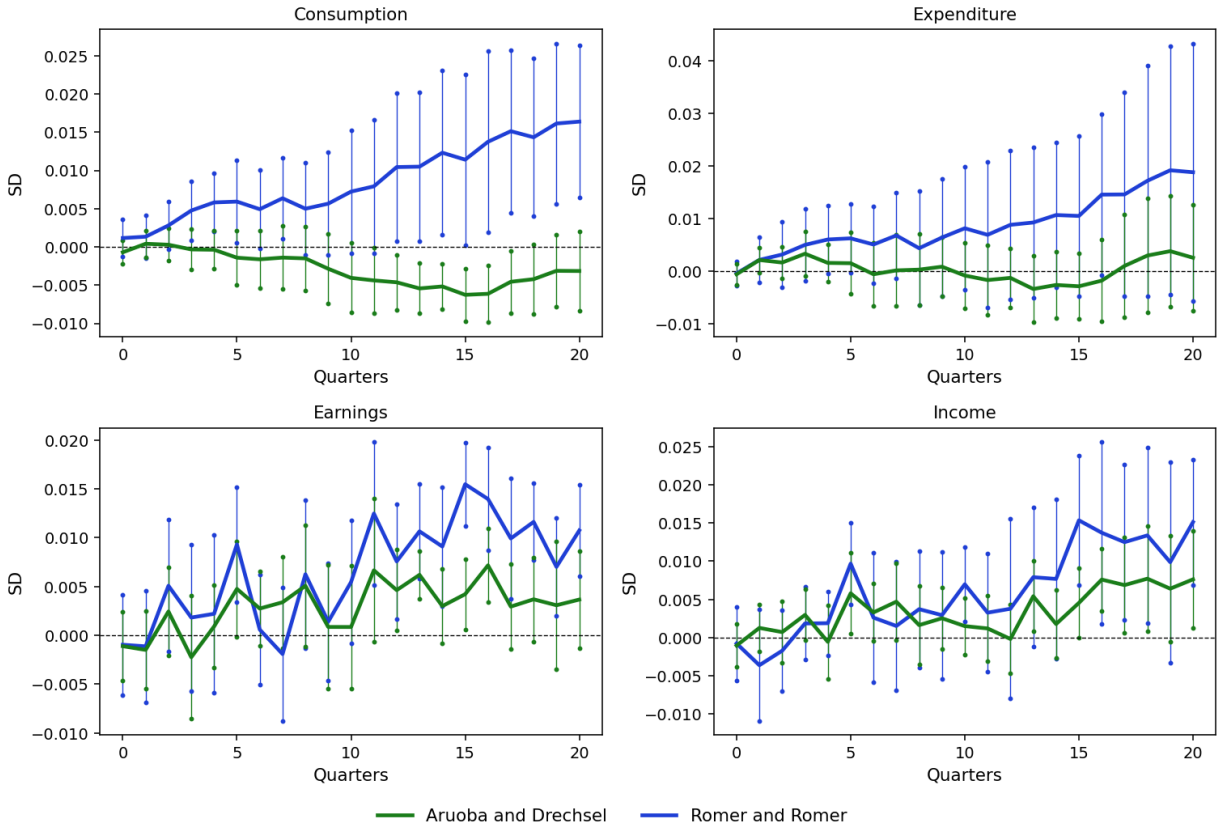
Note: This figure shows the impact of different monetary policy shocks on inequality with associated 1 standard deviation confidence interval using the local projection method. The inequality is measured by the Gini coefficient. The shocks are Aruoba and Drechsel (2024)(green), Romer and Romer (2004) (blue). The top left is consumption, the top right is expenditure, the bottom left is salary earnings, and the bottom right is total income. Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) starts from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.5: Different Monetary Policy Shocks on Inequality: 90-10



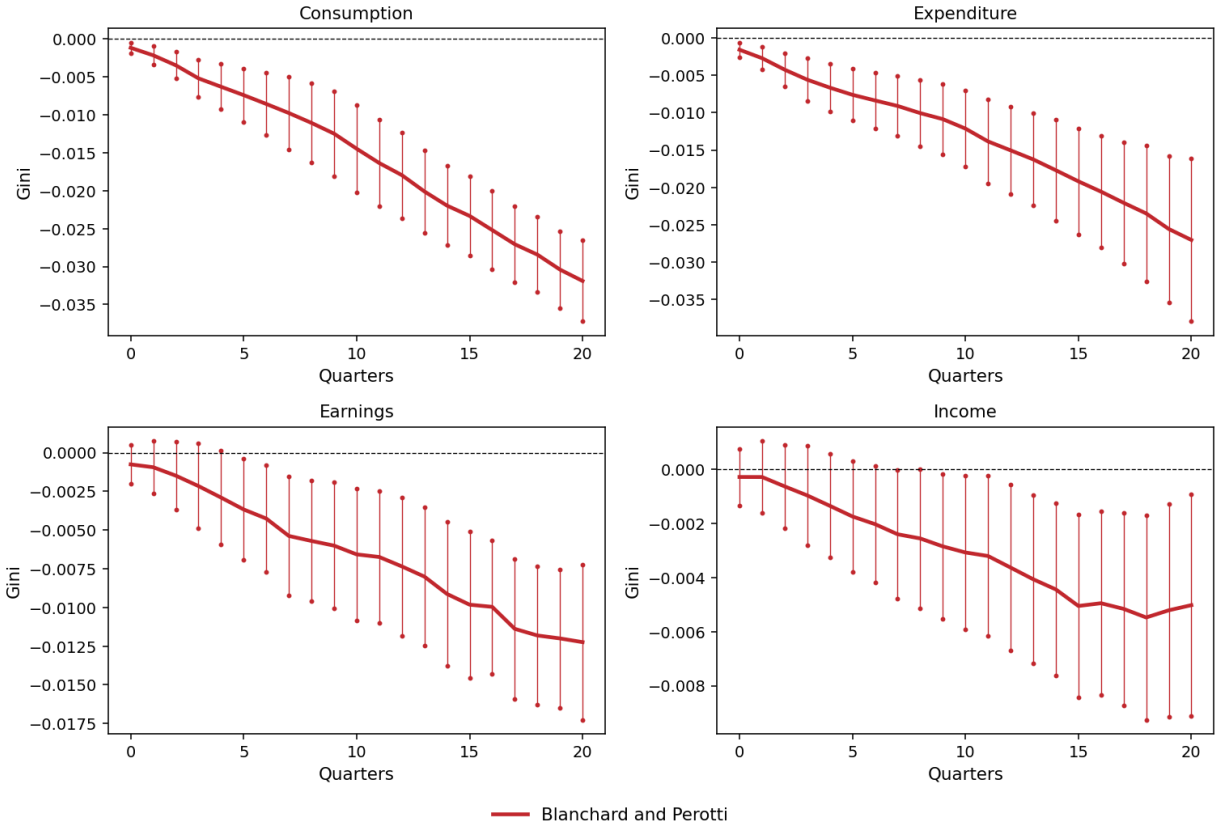
Note: This figure shows the impact of different monetary policy shocks on inequality with associated 90% confidence interval using the local projection method. The inequality is measured by the difference between the 90th percentile and the 10th percentile of the log levels in each distribution. The shocks are Aruoba and Drechsel (2024)(green), Romer and Romer (2004) (blue). The top left is consumption, the top right is expenditure, the bottom left is salary earnings, and the bottom right is total income. Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) starts from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.6: Different Monetary Policy Shocks on Inequality: Standard Deviation



Note: This figure shows the impact of different monetary policy shocks on inequality with associated 90% confidence interval using the local projection method. The inequality is measured by the cross-sectional standard deviation of the log value. The shocks are Aruoba and Drechsel (2024)(green), Romer and Romer (2004) (blue). The top left is consumption, the top right is expenditure, the bottom left is salary earnings, and the bottom right is total income. Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) starts from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.7: Fiscal Policy Shock on Inequality



Note: This figure shows the impact of fiscal policy shocks in Blanchard and Perotti (2002) on inequality with associated 90% confidence interval using the local projection method. The inequality is measured by the cross-sectional standard deviation of the log value (top row), Gini coefficient (middle row), and the difference between the 90th percentile and the 10th percentile of the log levels in each distribution (bottom). The columns from left to right are total income, salary earnings, expenditure, and consumption. The data period from 1980/01 to 2008/12.

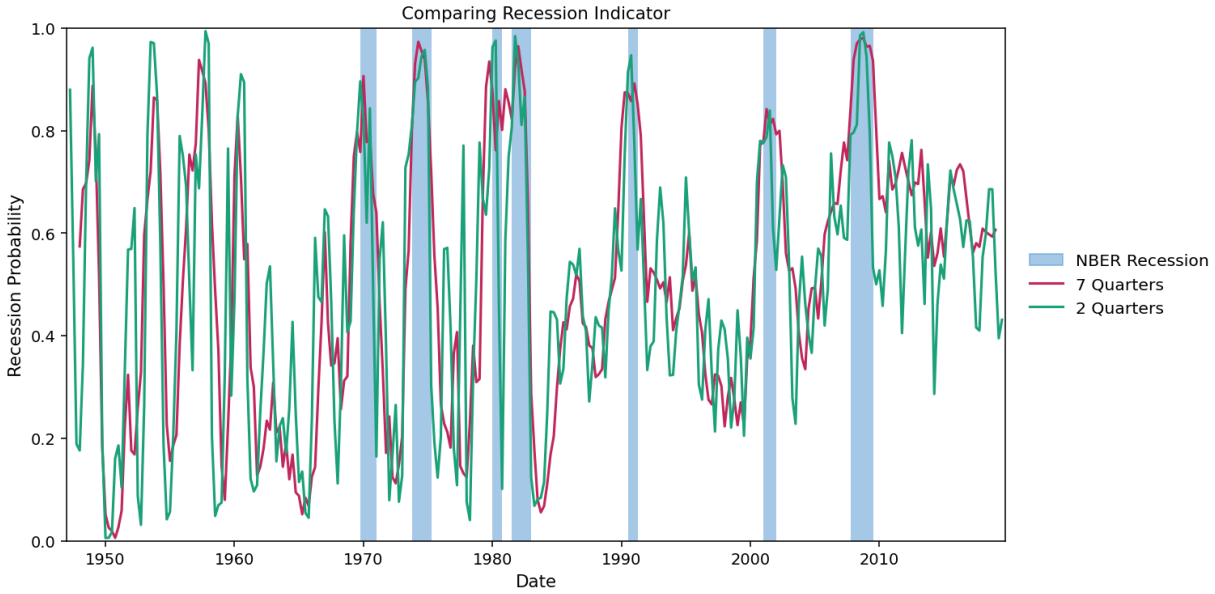
Here the dependent variable is the cumulated change in the inequality measure between $t-1$ and $t+h$, and η_t is the policy shock. As in the main text, the specification follows Coibion et al. (2017) and includes $I = 20$ state-interacted shock lags with $J = 2$ lags of the dependent variable, which the smooth-transition state and the stacked-panel estimator can support, for the Romer and Romer (2004) monetary policy shock and the fiscal policy shock. For the Aruoba and Drechsel (2024) monetary policy shock, due to the shorter available sample, we set $I = 15$ to be as close as possible. The alternative specification of the online Appendix A.4 (NBER dating with the contemporaneous shock only, Eq. (A.6)) is the counterpart used for the extended samples and the alternative shock series.

Recession Indicator The smooth-transition recession probability indicator of Auerbach and Gorodnichenko (2012b) used in the main text is constructed by calculating a 7-quarter average of real GDP growth, denoted as $z_t = (\log(GDP_{t+3}) - \log(GDP_{t-4}))/4$, and normalizing it by subtracting the mean and dividing by the standard deviation, yielding z_t^n. The recession indicator $F(z_t)$ is then computed using the logistic function $F(z_t) = \exp(\theta_z z_t^n)/(1 + \exp(\theta_z z_t^n))$, with the smoothing parameter θ_z set to 1.5, consistent with Auerbach and Gorodnichenko (2012b). Its advantage over a recession dummy is that it exploits continuous variation in the cyclical position, which is useful when the sample contains few discrete recessions.

These results are reported on the 1980-2008 sample, where the growth-based indicator is well behaved. We do not use it on the extended sample: because trend GDP growth declined persistently after 2008, a state variable built from demeaned GDP growth increasingly assigns high recession probabilities to normal post-2008 quarters, blurring the recession/expansion distinction that identifies the state-dependent coefficients; the extended-sample results therefore use NBER dating (online Appendix A.4). For robustness, we also consider a 2-quarter average instead of the 7-quarter average. We compare these indicators with NBER recession dates in Figure A.8; over 1980-2008 they align closely with the NBER chronology.

An Unemployment-Slack State Dating states by unemployment slack rather than by contractions, using the $\mathbb{1}\{U_{t-1} > 6.5\}$ indicator of Ramey and Zubairy (2018), reverses the sign of the consumption-Gini state difference (Romer and Romer (2004) -0.007 at three years; Ramey and Zubairy (2018) news $+0.013$). This is informative about the state rather than the fact: the slack indicator is dominated by high-unemployment *recovery* years (51 of its 63 slack quarters on 1980-2015 fall outside NBER contractions), so it contrasts depressed *levels* of activity with the rest of

Figure A.8: Comparing Recession Indicator



Note: This figure plots the time series of recession indicators using 7 quarters (red line) and 2 quarters (green line) averages compared to the NBER recessions (blue bar).

the sample rather than *falling* activity with rising, and our asymmetry is specific to the latter—consistent with the model's reference-dependence and with the generic below-trend null (online Appendix A.6).

Testing State-Shock Independence Following Gonçalves et al. (2024), we verify a key identification assumption for state-dependent analysis: policy shocks should not systematically influence the economy's state. This assumption could be violated if, for instance, a large interest rate hike or substantial government spending directly triggers a recession. To formally test this assumption, we estimate:

$$F(z_t) = \alpha + \beta^1 \eta_t + \beta^2 \eta_{t-1} + \beta^3 \eta_{t-2} + \varepsilon_t \tag{A.5}$$

where $F(z_t)$ is the recession indicator and η_t represents the policy shock at time t. We employ Newey-West standard errors with 20 lags to account for heteroskedasticity and serial correlation. The null hypothesis of no state-shock relationship implies $\beta^1 = \beta^2 = \beta^3 = 0$; a small F-statistic (large p-value) thus indicates that the shock and its recent lags do not move the probability of being in a recession, so the recession/expansion split in Eq. (1) is not driven by the policy shock itself.

Table A.1 presents our test results across different policy measures. Several findings support the validity of our identification strategy. First, most individual coefficients are statistically

Table A.1: Impact of policy shock on states

	AD	**RR**	**BP**
Shock_t	-0.366	0.000	0.081
	(0.295)	(0.001)	(0.126)
Shock_{t-1}	0.253	0.000	-0.073
	(0.211)	(0.001)	(0.082)
Shock_{t-2}	0.433	0.001*	-0.002
	(0.284)	(0.000)	(0.123)
Constant	0.510***	0.489***	0.382***
	(0.063)	(0.040)	(0.122)
F-statistic	1.959	1.904	0.727
p-value	0.129	0.131	0.537
N	71	153	153

Note: This table shows the impact of different policy shocks on the likelihood of reaching recessions. Policy shocks from left to right are: monetary shock in Aruoba and Drechsel (2024) (AD), monetary shock in Romer and Romer (2004) (RR), and fiscal policy shock (BP). Standard errors adjusted with Newey-West are in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. Data period varies in starting time subject to availability of data series.

insignificant, with only one exception (the second lag of RR showing marginal significance at 10%). Second, the F-tests fail to reject the null hypothesis of no joint significance for all specifications, with p-values ranging from 0.129 to 0.537. Third, the economic magnitude of the effects is small—for instance, a one-standard-deviation monetary policy shock in AD measure is associated with only a 0.366 percentage point change in recession probability. Together, these results suggest that our policy shocks are sufficiently modest to not systematically alter the state of the economy, validating our empirical approach.

Result: State Dependence Impact of Policy Shock on Inequality We present the state-dependent impacts of policy on inequality, with Figures A.9 and A.10 highlighting the effects on consumption and earnings inequality. Regardless of the shock source, contractionary monetary policy consistently increases consumption inequality during recessions, while expansionary fiscal policy reduces it. During expansions, these effects on consumption inequality are notably muted as shown in Figures A.9 and A.10. Similar patterns are observed for expenditure inequality, though with smaller effects, likely due to the inclusion of durable goods, which are less sensitive to short-term policy changes.

These observed patterns can be explained by two potential mechanisms. First, lower-income households might be disproportionately affected during recessions, experiencing larger wage reductions that widen consumption inequality when contractionary monetary policy is implemented, a channel that would operate through the earnings distribution. Alternatively, and as our state-dependent effective-curvature mechanism generates, the asymmetry can arise on the *consumption* margin with earnings roughly unchanged: in the loss regime the effective curvature of both types falls, but it falls *least* for the more loss-averse high-income type, so high-income households hold consumption nearest their reference point while low-income households, who rely on wages and experience the larger curvature decline, absorb the deeper contraction, raising the low-income consumption-output loading ($\chi^r > \chi^e$). The behavioral reading (consistent with Bauer and Swanson (2023) and Guvenen et al. (2021)) is thus an *incidence* statement about who bears the contraction, not a claim that high-income households cut spending or save more in absolute terms.

To distinguish between these mechanisms, we examine earnings inequality responses. If the first theory were correct, we would expect to see earnings inequality widen substantially during recessions in response to monetary policy shocks. However, as shown in the lower row of Figures A.9 and A.11, our results reveal no robust state dependence in the response of earnings

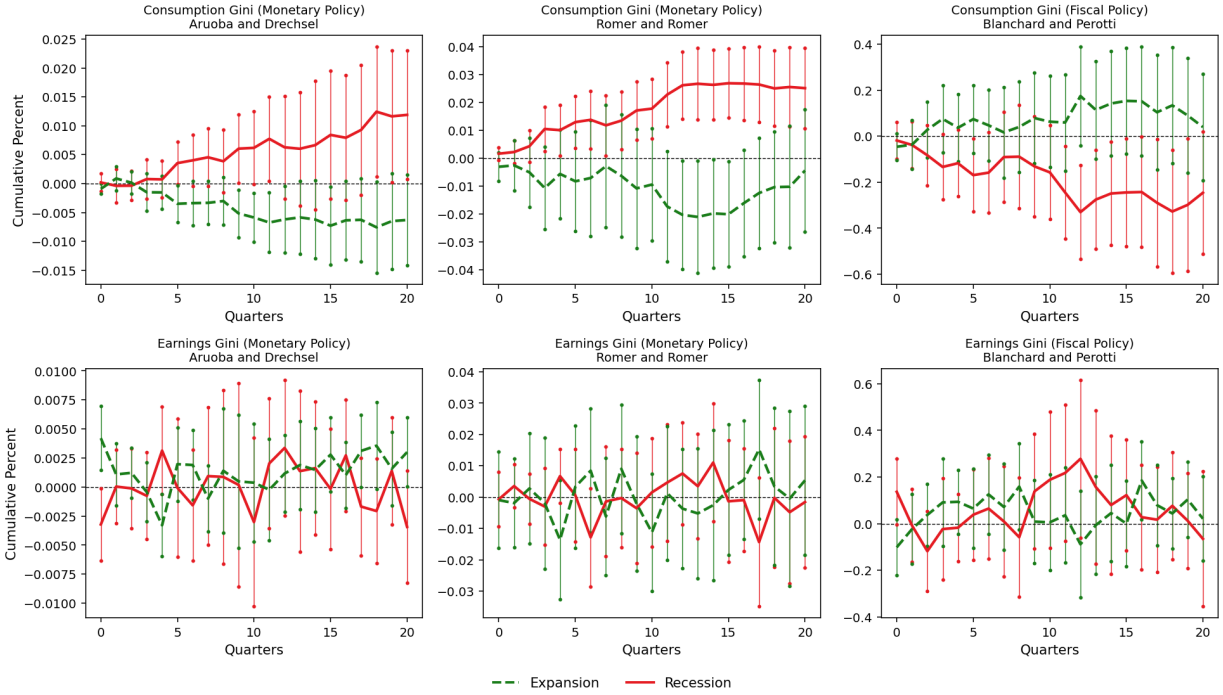
inequality to policy interventions: the estimates oscillate around zero with no consistent sign across state definitions rather than displaying the consistent asymmetry found for consumption. This finding aligns with our theoretical model's prediction that the asymmetric effects on consumption inequality are driven primarily by behavioral responses rather than changes in income distributions.

Interestingly, when examining broader income measures that include investment returns and savings income in Figures A.10 and A.12, we do observe some evidence of widening inequality under contractionary monetary policy and narrowing under expansionary fiscal policy during recessions. This pattern is consistent with our theoretical framework, in which the low-income consumption-output loading rises in downturns through the loss-aversion (effective-curvature) channel (high-income households' curvature falls least, so they hold consumption nearest their reference while the loading on low-income consumption steepens) widening the consumption and broad-income gap even as labor earnings remain relatively stable.

Further evidence supporting our loss aversion mechanism is visible in the standard deviation measures shown in Figures A.13 and A.14. Here, we observe clear divergence between consumption and earnings responses during recessions, with consumption inequality measures showing stronger state-dependent patterns than earnings measures.

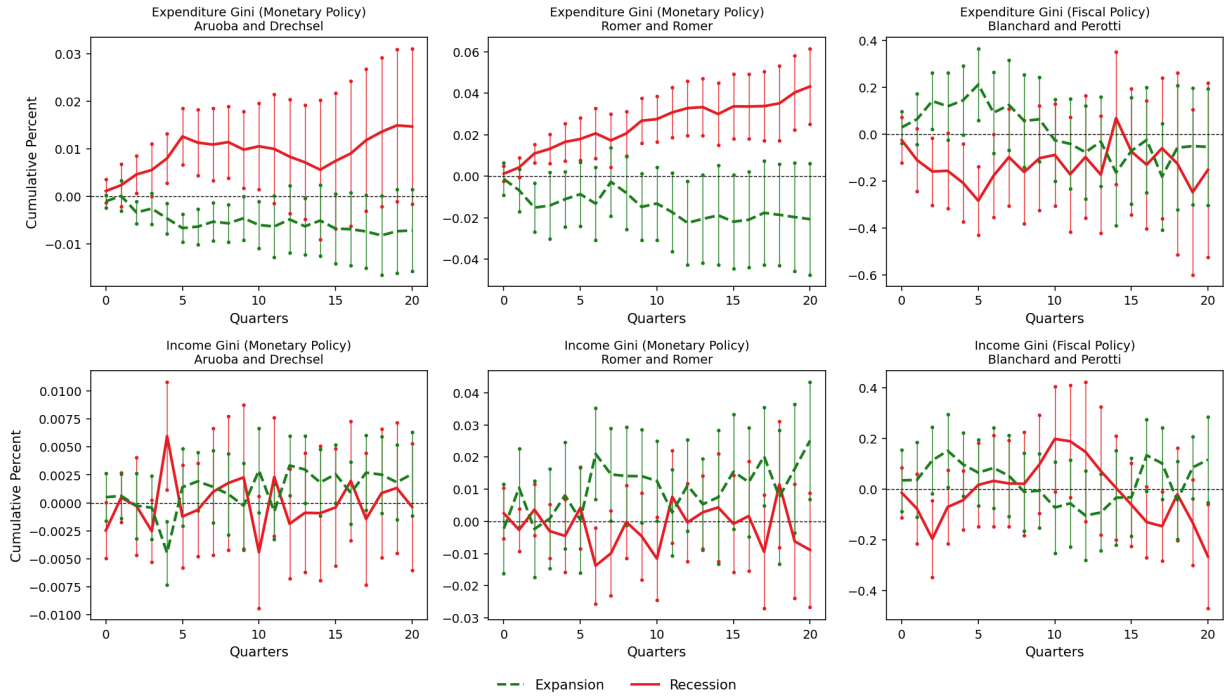
Overall, these results support our theoretical model's central premise: during recessions the effective curvature of both types falls, but least for the more loss-averse high-income type, so high-income households hold consumption nearest their reference point while low-income households, wage-reliant and experiencing the larger curvature decline, bear the deeper contraction ($\chi^r > \chi^e$), widening consumption and total-income inequality while leaving labor earnings inequality little changed. This incidence pattern aligns with our heterogeneous loss aversion mechanism, whose force is operative precisely in downturns, when consumption sits below its reference level.

Figure A.9: Different Policy Shocks on Consumption and Earnings Inequality in Business Cycles: Gini Coefficient



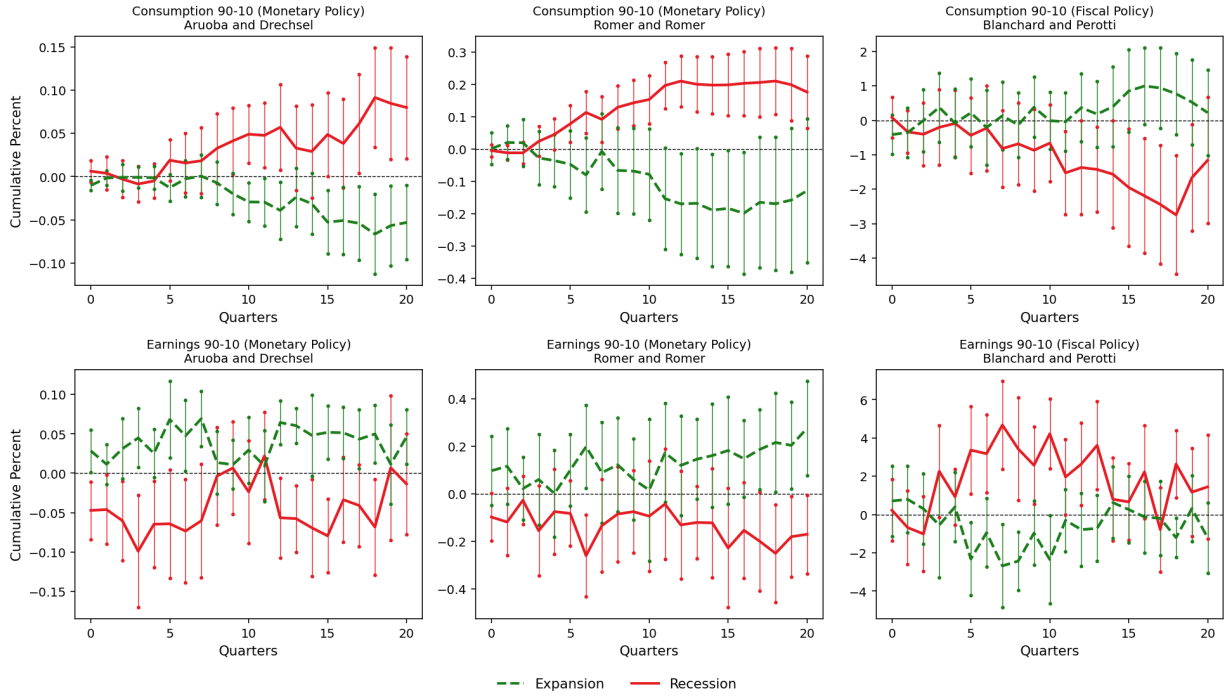
Note: This figure shows the impact of increasing 1 standard deviation different policy shocks on inequality in recession (red line) and expansion (green line) with associated 1 standard deviation confidence interval using state-dependence local projection method using 7-quarter average recession indicator. The inequality is measured by the consumption (upper row) and earnings (lower row) Gini coefficient. The shocks from left to right are: Aruoba and Drechsel (2024), Romer and Romer (2004) and fiscal shock in Blanchard and Perotti (2002). Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.10: Different Policy Shocks on Expenditure and Income Inequality in Business Cycles: Gini Coefficient



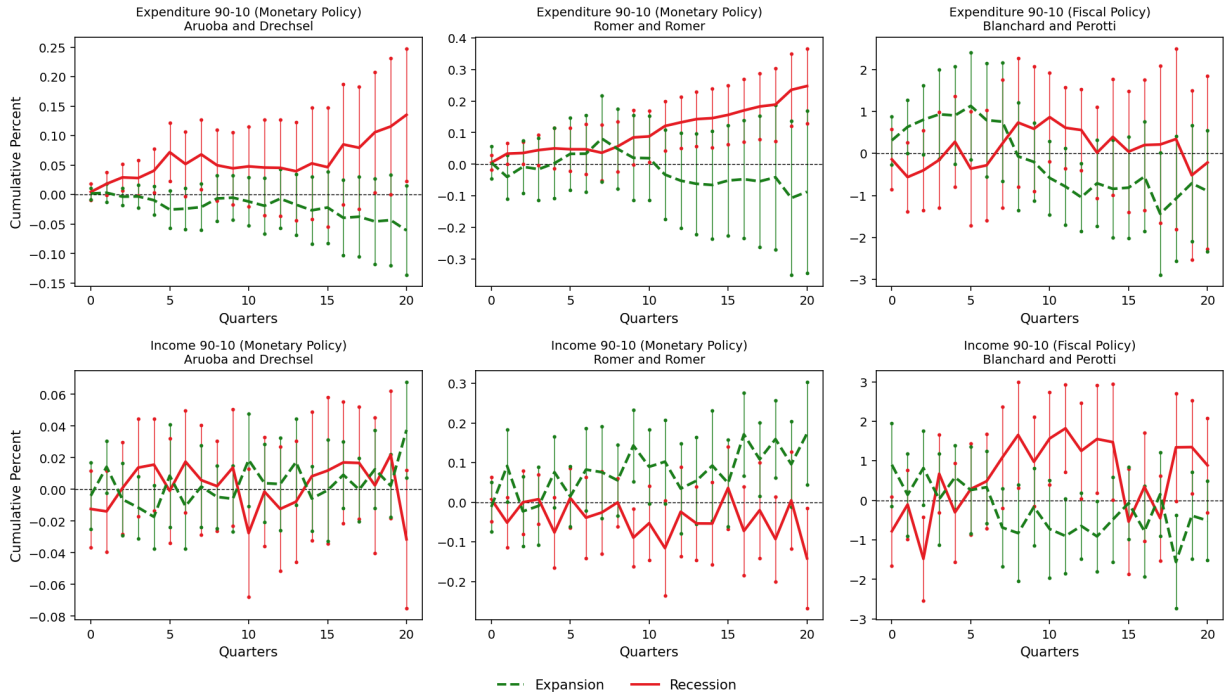
Note: This figure shows the impact of increasing 1 standard deviation different policy shocks on inequality in recession (red line) and expansion (green line) with associated 1 standard deviation confidence interval using state-dependence local projection method using 7-quarter average recession indicator. The inequality is measured by the expenditure (upper row) and income (lower row) Gini coefficient. The shocks from left to right are: Aruoba and Drechsel (2024), Romer and Romer (2004) and fiscal shock in Blanchard and Perotti (2002). Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.11: Different Policy Shocks on Consumption and Earnings Inequality in Business Cycles: 90-10



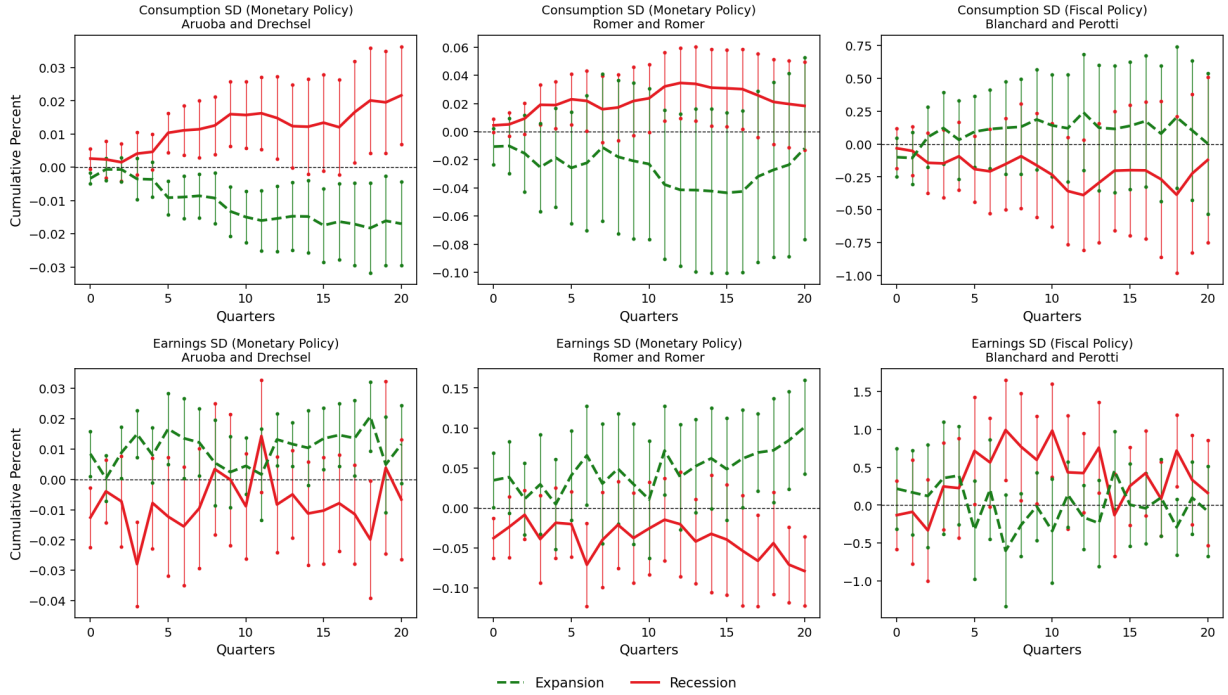
Note: This figure shows the impact of increasing 1 standard deviation different policy shocks on inequality in recession (red line) and expansion (green line) with associated 1 standard deviation confidence interval using state-dependence local projection method using 7-quarter average recession indicator. The inequality is measured by the difference between the 90th percentile and the 10th percentile of the log consumption (upper row) and earnings (lower row) in each distribution. The shocks from left to right are: Aruoba and Drechsel (2024), Romer and Romer (2004) and fiscal shock in Blanchard and Perotti (2002). Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.12: Different Policy Shocks on Expenditure and Income Inequality in Business Cycles: 90-10



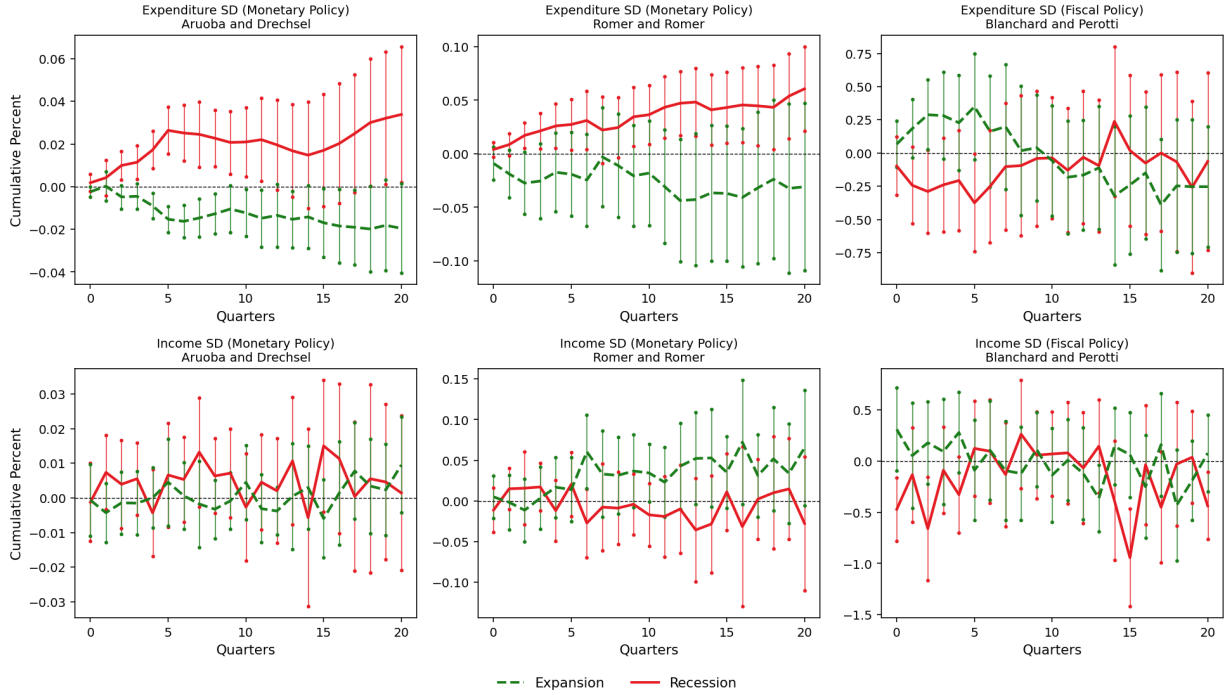
Note: This figure shows the impact of increasing 1 standard deviation different policy shocks on inequality in recession (red line) and expansion (green line) with associated 1 standard deviation confidence interval using state-dependence local projection method using 7-quarter average recession indicator. The inequality is measured by the difference between the 90th percentile and the 10th percentile of the log expenditure (upper row) and income (lower row) in each distribution. The shocks from left to right are: Aruoba and Drechsel (2024), Romer and Romer (2004) and fiscal shock in Blanchard and Perotti (2002). Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Figure A.13: Different Policy Shocks on Consumption and Earnings Inequality in Business Cycles: Standard Deviation



Note: This figure shows the impact of increasing 1 standard deviation different policy shocks on inequality in recession (red line) and expansion (green line) with associated 1 standard deviation confidence interval using state-dependence local projection method using 7-quarter average recession indicator. The inequality is measured by the cross-sectional standard deviation of the log consumption (upper row) and log earnings (lower row). The shocks from left to right are: Aruoba and Drechsel (2024), Romer and Romer (2004) and fiscal shock in Blanchard and Perotti (2002). Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

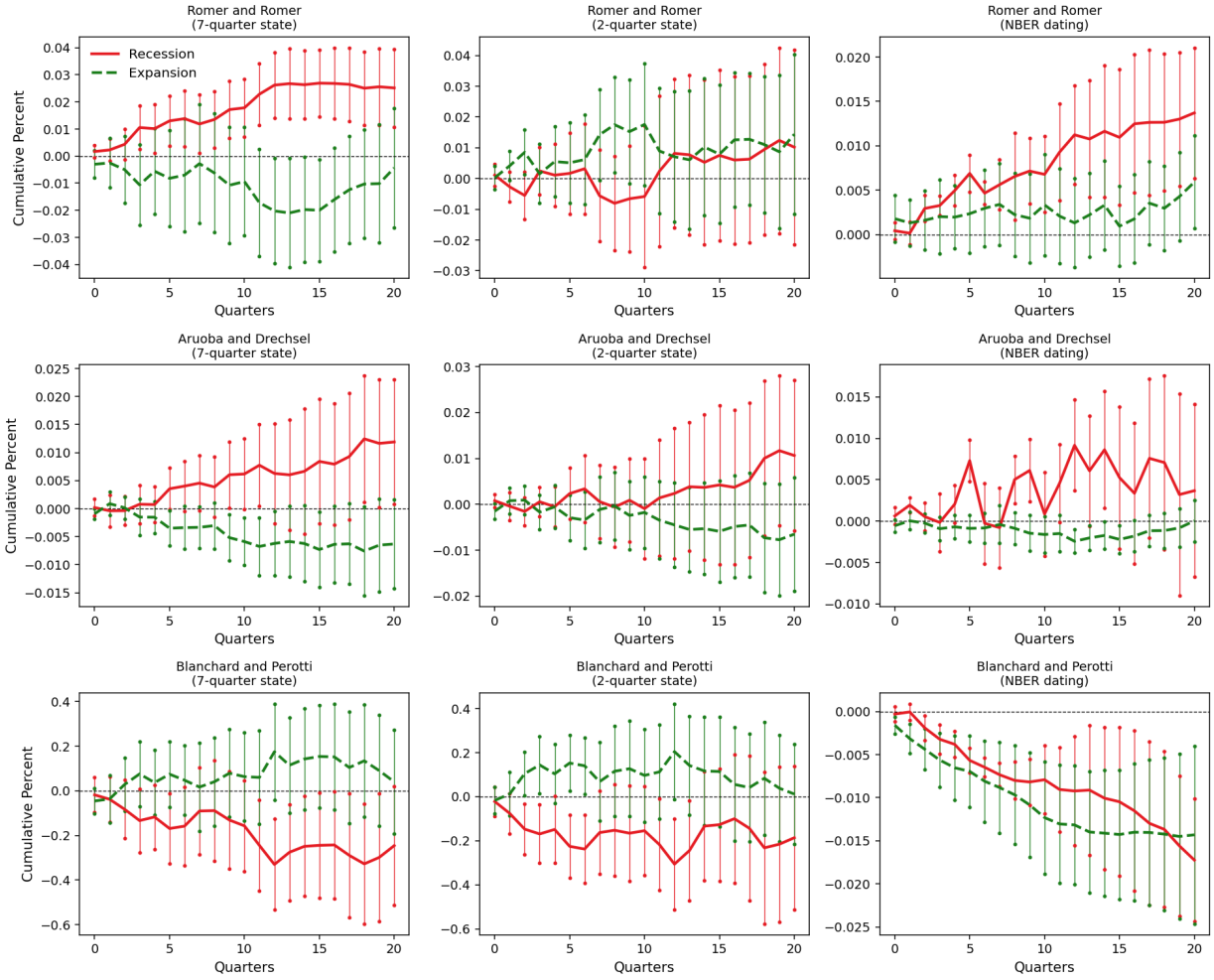
Figure A.14: Different Policy Shocks on Consumption and Earnings Inequality in Business Cycles: Standard Deviation



Note: This figure shows the impact of increasing 1 standard deviation different policy shocks on inequality in recession (red line) and expansion (green line) with associated 1 standard deviation confidence interval using state-dependence local projection method using 7-quarter average recession indicator. The inequality is measured by the cross-sectional standard deviation of the log expenditure (upper row) and log income (lower row). The shocks from left to right are: Aruoba and Drechsel (2024), Romer and Romer (2004) and fiscal shock in Blanchard and Perotti (2002). Data period varies in starting time subject to availability of data series: Aruoba and Drechsel (2024) start from 1982/10, Romer and Romer (2004) start from 1980/01 but all end at 2008/12.

Robustness to the recession dating. The 7-quarter smooth-transition state above is one of several ways to date recessions. Figure A.15 plots the recession and expansion consumption-Gini responses for the three headline shocks under all three datings used in the paper: the 7-quarter smooth transition, the 2-quarter version of Santoro et al. (2014), and NBER recession dates. The asymmetry survives every dating. The monetary recession response (Romer and Romer (2004), Aruoba and Drechsel (2024)) lies above the expansion response in each column, and the fiscal ordering (Blanchard and Perotti (2002)) reverses in each column; the magnitude is largest under the 7-quarter state and smaller under the 2-quarter and NBER datings, but the sign and state-ordering do not change.

Figure A.15: Robustness to the Recession Dating: Consumption-Gini Responses under Three State Definitions



Note: Cumulated consumption-Gini responses in recessions (solid red) and expansions (dashed green) to a one-standard-deviation shock, for the three headline shocks (rows: Romer and Romer (2004), Aruoba and Drechsel (2024), Blanchard and Perotti (2002)) under three recession datings (columns: the 7-quarter smooth-transition state, the 2-quarter state of Santoro et al. (2014), and NBER recession dates). The 7- and 2-quarter columns use the joint Driscoll-Kraay estimator of Eq. (1); the NBER column uses the per-horizon estimator of Eq. (A.6). Vertical bars are 90% confidence intervals.

A.4 Alternative Shock Series and NBER Recession Dating

This appendix re-establishes the main-text facts under a fully alternative specification: NBER recession dating, alternative monetary and fiscal shock identifications, samples extended to 2019 where the identifications allow, Newey-West inference horizon by horizon (online Appendix A.6). The estimating equation is the per-horizon variant of the main-text specification,

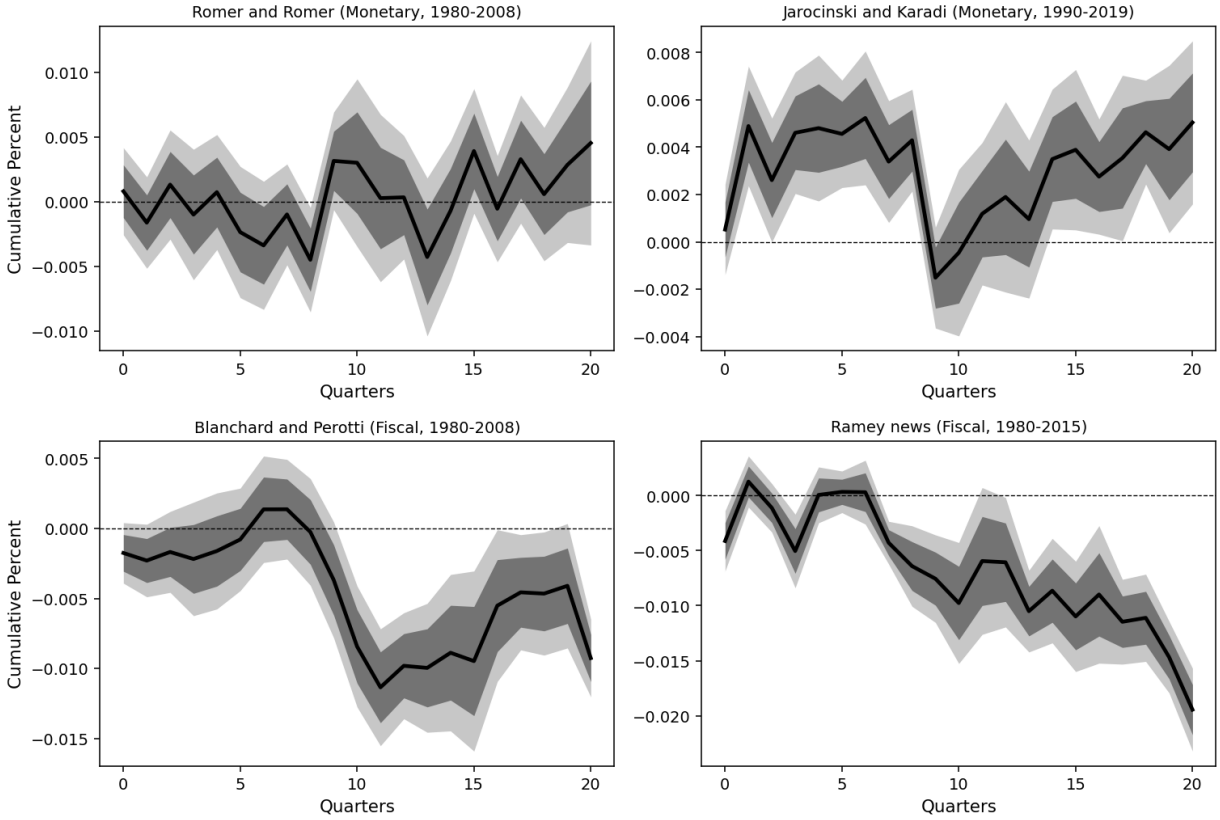
$$
\begin{aligned}
x_{t+h} - x_{t-1} = F(z_{t-1}) & \left[c^{r,(h)} + (\beta^r)^{(h)} \eta_t + \sum_{j=1}^{J} (\alpha_j^r)^{(h)} x_{t-j} \right] \\
& + (1 - F(z_{t-1})) \left[c^{e,(h)} + (\beta^e)^{(h)} \eta_t + \sum_{j=1}^{J} (\alpha_j^e)^{(h)} x_{t-j} \right] + \varepsilon_{t+h}, \\
& h = 0, \ldots, H.
\end{aligned}
\tag{A.6}
$$

where the dependent variable $x_{t+h} - x_{t-1}$ is the *cumulated change* in the outcome between $t-1$ and horizon $t+h$, so the contemporaneous-shock coefficients $(\beta^r)^{(h)}$ and $(\beta^e)^{(h)}$ trace out cumulated impulse responses directly. Three specification choices deserve statement: each business-cycle bracket carries its *own* intercept and its own $J = 2$ lags of the dependent variable, so the regression is a full sample split; the state indicator is *lagged one quarter*, $F(z_{t-1})$, so the state is measured before the shock arrives, which removes the within-quarter feedback that Gonçalves et al. (2024) warn about; and only the *contemporaneous* shock enters each bracket, following the horizon-by-horizon convention of Jordà (2005) and Ramey and Zubairy (2018). We set $H = 20$. Here $F(z_{t-1})$ is the NBER recession dummy, lagged one quarter; the NBER dating is not mechanically tied to GDP growth relative to trend, which matters because trend growth declined markedly after 2008, so growth-based smooth-transition states increasingly misclassify normal post-2008 quarters as high-recession-probability states on extended samples (online Appendix A.3). Inference is by Newey and West (1987) standard errors horizon by horizon on the cumulated dependent variable (truncation lag $h+1$). The alternative headline series are the high-frequency monetary shock of Jarociński and Karadi (2020) (1990-2019), which separates genuine surprises from central-bank information effects, and the narrative military spending news variable of Ramey and Zubairy (2018) (1980-2015) for fiscal policy; the Aruoba and Drechsel (2024) series covers the pre-1990 window excluding the Volcker disinflation, and the online Appendix A.5 extends the monetary analysis across identification families.

These results are shown as the recession-minus-expansion difference $\beta^r - \beta^e$, the *state difference* of the Gini. The consumption difference under this alternative specification is Figure 4 in the main text; its earnings counterpart is Figure A.16 below, and Figure A.17 decomposes the consumption difference into the two states separately, $(\beta^r)^{(h)}$ and $(\beta^e)^{(h)}$.

Monetary policy. Figure A.17 decomposes the monetary state difference into the two states. The asymmetry is consistent across identification families—narrative and NLP-based (Romer and Romer (2004), Aruoba and Drechsel (2024)) and high-frequency (Jarociński and Karadi (2020))—

Figure A.16: Robustness Across Identifications: Earnings-Inequality State Difference under NBER Dating



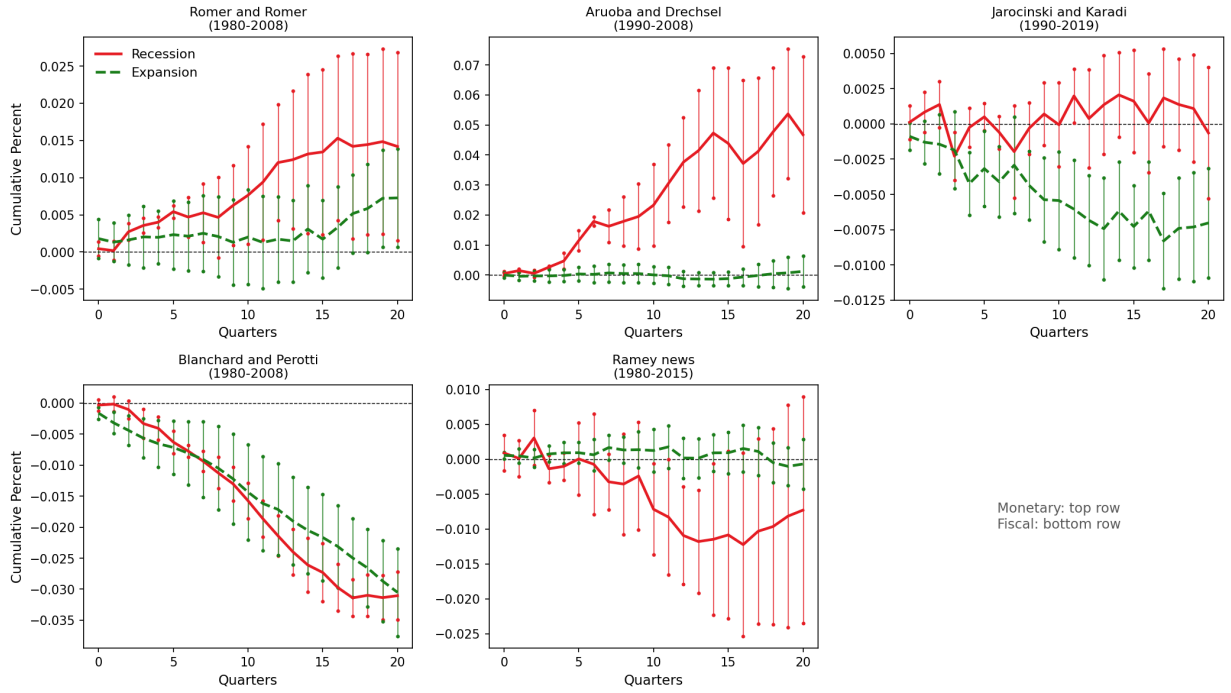
Note: As in Figure 4, for the *earnings* Gini: the same identifications, samples, dating, and estimator (Eq. (A.6)), with the earnings series spliced through 2019. Dark and light shaded bands are 68% and 90% Newey-West intervals (truncation lag $h+1$).

with the same sign and state-ordering on every series and magnitudes that track the size of the surprise; the strength of the evidence is this consistency rather than any single specification. Per-series estimates are in the online Appendix A.6 (Table A.2).

Fiscal policy. The bottom row of Figure 4 shows the fiscal result. Under the Ramey and Zubairy (2018) news identification, expansionary fiscal shocks reduce consumption inequality by more in recessions (three-year state difference -0.011 over 1980-2015 and -0.017 ending in 2008; Table A.2), stable across episode deletions. The fiscal asymmetry provides supporting evidence alongside the monetary fact, with more moderate magnitudes.

Earnings inequality. Figure A.16 makes the wedge visible under the alternative specification: the monetary earnings state differences are small, and none reproduces the positive monetary

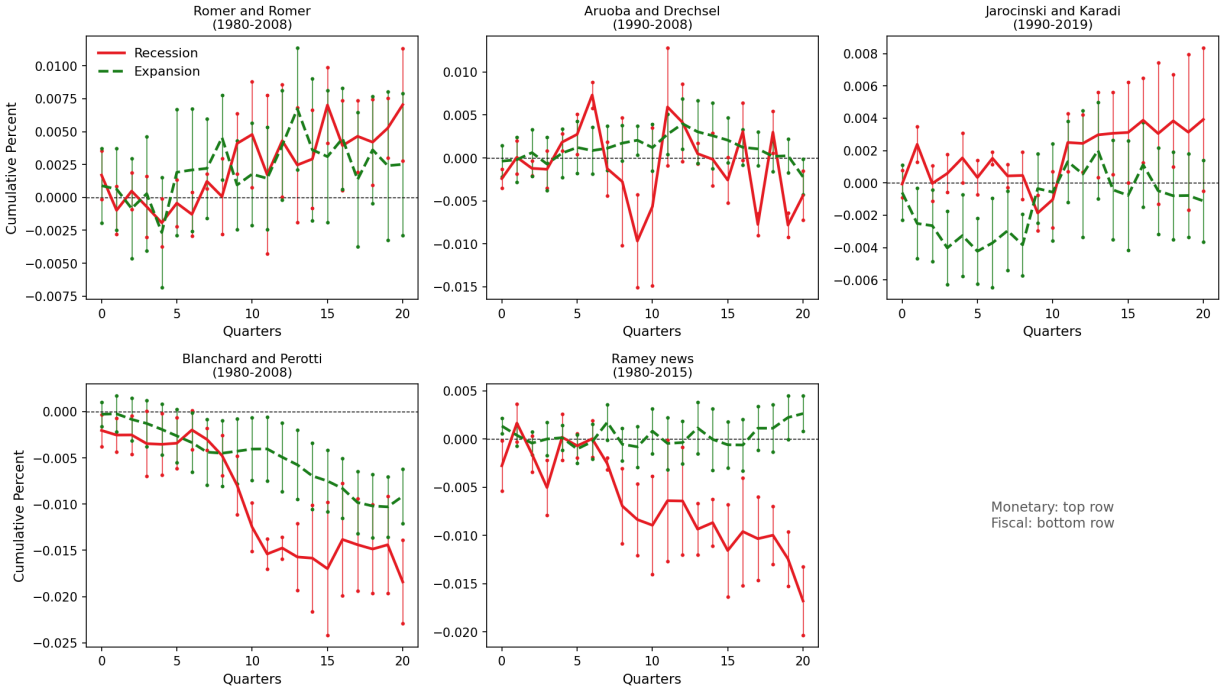
Figure A.17: Recession versus Expansion Responses of Consumption Inequality across Identifications



Note: The consumption-Gini difference of Figure 4 decomposed into the two states separately, for five identifications, all under NBER dating (Eq. (A.6), per-horizon Newey-West; vertical bars are 90% confidence intervals). Top row, monetary policy: Romer and Romer (2004) (1980-2008), Aruoba and Drechsel (2024) (1990-2008), Jarociński and Karadi (2020) (1990-2019). Bottom row, fiscal policy: Blanchard and Perotti (2002) (1980-2008), Ramey and Zubairy (2018) news (1980-2015). Each panel is a single shock on its own sample; the gap between the two lines is the state difference. All the alternative monetary series are decomposed in the online Appendix A.5.

consumption asymmetry, so the earnings null holds under the alternative identifications and dating too. Figure A.18 decomposes the earnings differences into the two states, mirroring Figure A.17: for the monetary shocks the recession and expansion responses nearly coincide. The fiscal earnings differences behave differently (Table A.2, Panel D): they are negative and in places significant, with magnitudes comparable to the fiscal consumption differences, indicating that a recession-timed spending expansion compresses earnings inequality directly, consistent with government spending operating in part through labor demand and the hours and participation margins that Heathcote et al. (2020) show drive cyclical earnings dispersion—a margin from which the model abstracts, since wages are common across household types. This is why the monetary fact carries the identifying weight for the consumption-side mechanism—there consumption inequality moves with no earnings counterpart—while the fiscal asymmetry, which partly runs through earnings, is read as supporting evidence, as in the main text.

Figure A.18: Recession versus Expansion Responses of Earnings Inequality across Identifications



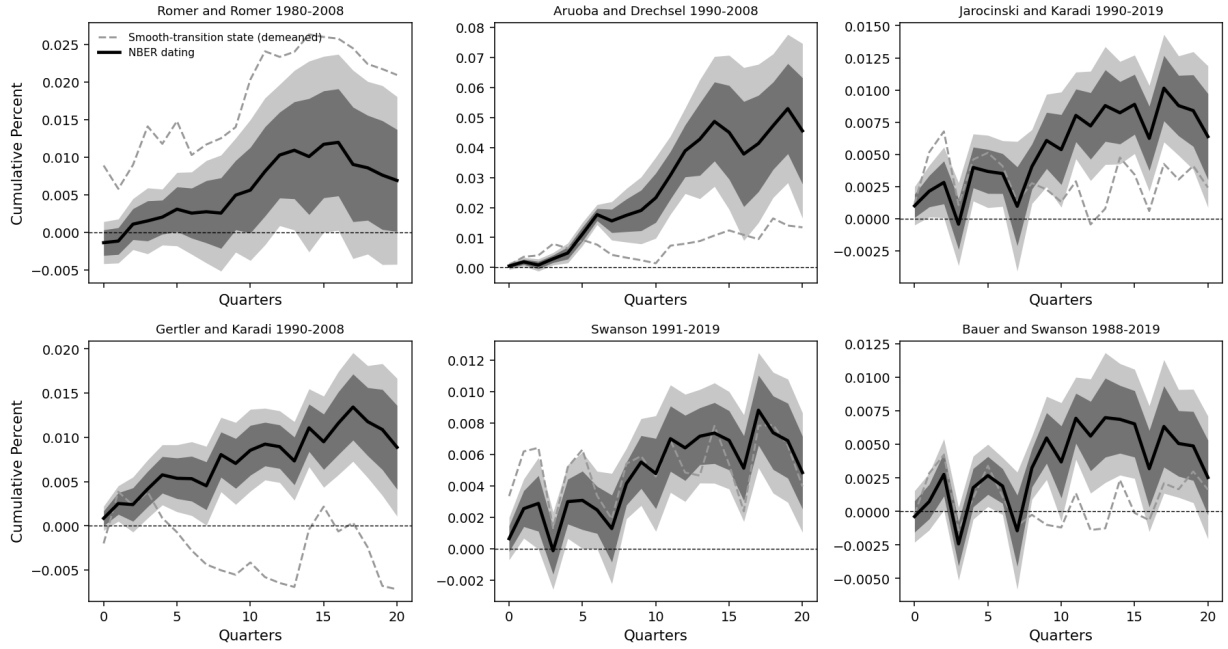
Note: The earnings-Gini state difference decomposed into the two states separately, for the same five identifications, samples, dating, and estimator as Figure A.17 (vertical bars are 90% confidence intervals; the earnings series is spliced through 2019). For the monetary shocks (top row) the recession and expansion responses nearly coincide; for the fiscal shocks (bottom row) the recession response is more negative, as in Panel D of Table A.2.

A.5 Alternative Monetary Shock Series

Our baseline evidence rests on the narrative identifications, which end in 2008 (online Appendix A.1). Figure A.19 shows the fact is not specific to them: it re-estimates the recession-minus-expansion difference of the consumption-Gini response, $\beta^r - \beta^e$, under NBER dating, across the main families of monetary identification—narrative, NLP-based, and high-frequency futures-market surprises, each on its own sample. The sign and state-ordering of the difference are unchanged across every series and sample; only the magnitude varies with the size of the underlying surprise. Table A.2 reports the corresponding point estimates.

Figure A.20 repeats the exercise for the earnings Gini on the same series and samples, using the earnings series spliced through 2019: none of the six series produces a state difference comparable to its consumption counterpart, so the earnings null of the main text holds across the same identification families.

Figure A.19: Alternative Monetary Shock Series: State Difference of the Consumption Gini Response



Note: Each panel plots the cumulated recession-minus-expansion difference, $\beta^r - \beta^e$, of the consumption Gini response to a one-standard-deviation contractionary monetary shock, estimated from Eq. (A.6), under NBER recession dating (solid black; dark and light shading are 68% and 90% Newey-West bands) and under the demeaned smooth-transition state variable (dashed grey). Shock series and samples as labeled in the panel titles.

A.6 Additional Robustness Checks

This appendix reports the full set of state-dependent estimates across shock series and samples. Estimates are from Eq. (A.6) under NBER dating, with Newey-West inference (truncation lag $h + 1$).

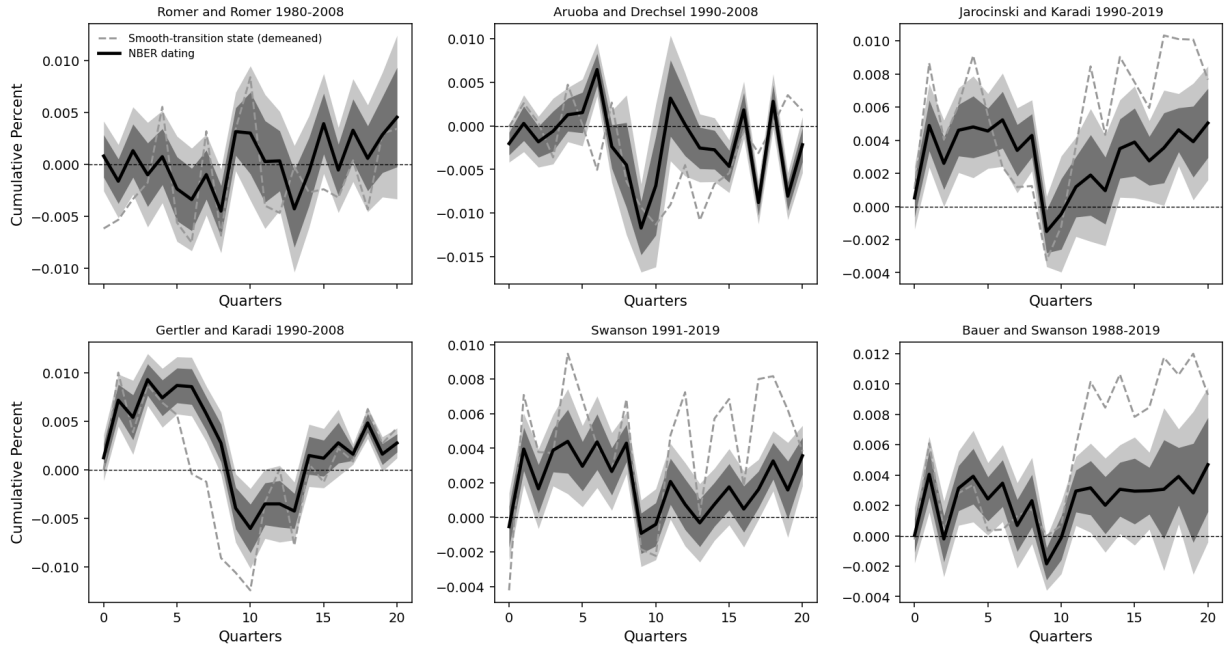
We also examined a state that classifies quarters by consumption below trend rather than by NBER dates: it yields no significant state differences for any shock series. As with the unemployment-slack state of Appendix A.3, this is informative about the state variable rather than about the fact—the asymmetry is specific to dated contractions, periods in which activity is falling, rather than to periods in which its level is depressed.

Table A.2: State Difference $\beta^r - \beta^e$ Across Monetary and Fiscal Shock Series

Shock	Sample	$h = 4$	$h = 8$	$h = 12$	$h = 20$
Panel A: Consumption Gini, contractionary monetary shock (Figure A.19)					
Romer and Romer	1980-2008	$+0.002$	$+0.003$	$+0.010^{*}$	$+0.007$
		(0.002)	(0.005)	(0.006)	(0.007)
Aruoba and Drechsel	1990-2008	$+0.005^{**}$	$+0.017^{***}$	$+0.039^{***}$	$+0.046^{**}$
		(0.002)	(0.005)	(0.009)	(0.018)
Jarocinski and Karadi	1990-2019	$+0.004^{**}$	$+0.004^{**}$	$+0.007^{***}$	$+0.006^{*}$
		(0.002)	(0.002)	(0.003)	(0.003)
Gertler and Karadi	1990-2008	$+0.006^{***}$	$+0.008^{***}$	$+0.009^{***}$	$+0.009^{*}$
		(0.002)	(0.003)	(0.002)	(0.005)
Swanson	1991-2019	$+0.003^{**}$	$+0.004^{***}$	$+0.006^{***}$	$+0.005^{**}$
		(0.001)	(0.001)	(0.002)	(0.002)
Bauer and Swanson	1988-2019	$+0.002$	$+0.003^{**}$	$+0.006^{**}$	$+0.003$
		(0.001)	(0.001)	(0.003)	(0.003)
Panel B: Consumption Gini, expansionary fiscal shock (Figure 4)					
Ramey news	1980-2015	-0.002	-0.005	-0.011^{**}	-0.007
		(0.002)	(0.005)	(0.005)	(0.010)
Ramey news	1980-2008	-0.004^{**}	-0.011^{***}	-0.017^{***}	-0.022^{**}
		(0.002)	(0.004)	(0.006)	(0.009)
Blanchard and Perotti	1980-2008	$+0.002$	-0.001	-0.004	-0.001
		(0.002)	(0.004)	(0.005)	(0.005)
Panel C: Earnings Gini, contractionary monetary shock (Figure A.20)					
Romer and Romer	1980-2008	$+0.001$	-0.004^{*}	$+0.000$	$+0.005$
		(0.003)	(0.002)	(0.003)	(0.005)
Aruoba and Drechsel	1990-2008	$+0.001$	-0.004	$+0.000$	-0.002
		(0.002)	(0.005)	(0.004)	(0.002)
Jarocinski and Karadi	1990-2019	$+0.005^{**}$	$+0.004^{***}$	$+0.002$	$+0.005^{**}$
		(0.002)	(0.001)	(0.002)	(0.002)
Gertler and Karadi	1990-2008	$+0.007^{***}$	$+0.003$	-0.003	$+0.003^{***}$
		(0.002)	(0.002)	(0.002)	(0.001)
Swanson	1991-2019	$+0.004^{**}$	$+0.004^{***}$	$+0.001$	$+0.004^{***}$
		(0.002)	(0.001)	(0.001)	(0.001)
Bauer and Swanson	1988-2019	$+0.004^{**}$	$+0.002$	$+0.003^{*}$	$+0.005$
		(0.002)	(0.002)	(0.002)	(0.003)
Panel D: Earnings Gini, expansionary fiscal shock					
Ramey news	1980-2015	$+0.000$	-0.006^{***}	-0.006^{*}	-0.019^{***}
		(0.002)	(0.002)	(0.004)	(0.002)
Ramey news	1980-2008	-0.001	-0.009^{***}	-0.006	-0.020^{**}
		(0.002)	(0.001)	(0.004)	(0.008)
Blanchard and Perotti	1980-2008	-0.002	-0.000	-0.010^{***}	-0.009^{***}
		(0.003)	(0.002)	(0.002)	(0.002)

Note: This table reports the statistical inference behind Figures A.19, A.20, and 4: cumulated recession-minus-expansion differences of the consumption (Panels A–B) and earnings (Panels C–D) Gini to a one-standard-deviation shock at the horizons of Table 2, estimated from Eq. (A.6) under NBER recession dating. The earnings series is spliced through 2019. Standard errors adjusted with Newey-West (truncation lag $h + 1$) are in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Figure A.20: Alternative Monetary Shock Series: State Difference of the Earnings Gini Response



Note: Each panel plots the cumulated recession-minus-expansion difference, $\beta^r - \beta^e$, of the earnings Gini response to a one-standard-deviation contractionary monetary shock, estimated from Eq. (A.6), under NBER recession dating (solid black; dark and light shading are 68% and 90% Newey-West bands) and under the demeaned smooth-transition state variable (dashed grey). Shock series and samples exactly as in Figure A.19; the earnings series is spliced through 2019; the results are essentially identical on the original series ending in 2008.

B Theoretical Proof

B.1 The Firm Block

Final-goods output aggregates intermediate varieties through the CES technology

$$C_t = \left[\int_0^1 C_t(z)^{\frac{\epsilon-1}{\epsilon}}\, dz\right]^{\frac{\epsilon}{\epsilon-1}}, \tag{B.1}$$

with $\epsilon > 1$. Cost minimization by the final-goods firm gives the demand curve and price index used in the main text,

$$C_t(z) = \left[\frac{P_t(z)}{P_t}\right]^{-\epsilon} C_t, \qquad P_t = \left[\int_0^1 P_t(z)^{1-\epsilon}\, dz\right]^{\frac{1}{1-\epsilon}}. \tag{B.2}$$

Intermediate firms are monopolistically competitive with linear technology $Y_t(z) = A_t N_t(z)$, where $A_t = \rho_A A_{t-1} + \epsilon_A$ with $\epsilon_A \sim N(0, \sigma_A^2)$, and reset prices with probability $1 - \Theta$ (Calvo, 1983). Under optimal pricing the desired relative price satisfies

$$\frac{P_t^*(z_t)}{P_t^*} = 1 = \frac{\epsilon W_t^*}{(1+\tau_s)(\epsilon - 1)}, \tag{B.3}$$

so the subsidy $\tau_s = (\epsilon - 1)^{-1}$, financed by the firm tax $T_t^F = \tau_s Y_t$, eliminates the markup and delivers zero net profit in steady state, $D_t = Y_t - W_t N_t = 0$.

B.2 Deriving the State-Dependence Consumption Response

The derivation follows the appendix of Pfäuti and Seyrich (2022). Using the first-order condition for labor in expansions, we obtain:

$$(N_t^L)^{\phi} = W_t[\alpha(C_t^L)^{-1} + (1-\alpha)(C_t^L)^{-\theta-1}]. \tag{B.4}$$

Similarly, we can obtain the conditions during the recession such that:

$$(N_t^L)^{\phi} = W_t[\alpha(C_t^L)^{-1} + (1-\alpha)(-\lambda^L)\frac{\exp\left(\frac{\theta}{\lambda^L}\ln\left(C_t^L\right)\right)}{C_t^L}], \tag{B.5}$$

and the conditions for the high-income household are analogous. Using the market clearing condition and the steady-state condition that all households have equal consumption, with $\zeta \in [e, r]$ as the index of expansion and recession, log-linearizing the labor condition gives:

$$\phi n_{\zeta,t}^L = w_{\zeta,t} - \sigma_\zeta^L c_{\zeta,t}^L, \tag{B.6}$$

$$\phi n_{\zeta,t}^H = w_{\zeta,t} - \sigma_\zeta^H c_{\zeta,t}^H, \tag{B.7}$$

where $\sigma_e^i = 1 + (1-\alpha)\theta, \sigma_r^i = 1 - (1-\alpha)\frac{\theta}{\lambda^i}$, as obtained by Santoro et al. (2014) without habit formation. In expansions the effective curvature is common across types ($\sigma_e^H = \sigma_e^L \equiv \sigma_e$); in recessions it differs across types ($\sigma_r^H > \sigma_r^L$ under $\lambda^H > \lambda^L$).

Since we assume both household types have the same productivity, and thus the same wage, the labor market condition implies:

$$\phi n_{\zeta,t}^L + \sigma_\zeta^L c_{\zeta,t}^L = \phi n_{\zeta,t}^H + \sigma_\zeta^H c_{\zeta,t}^H. \tag{B.8}$$

With the market clearing conditions ($n_{\zeta,t} = h n_{\zeta,t}^H + (1-h) n_{\zeta,t}^L$, $c_{\zeta,t} = h c_{\zeta,t}^H + (1-h) c_{\zeta,t}^L$) and using $y_{\zeta,t} = c_{\zeta,t} = n_{\zeta,t}$, after rearrangement we obtain:

$$\phi n_{\zeta,t}^L + \sigma_\zeta^L c_{\zeta,t}^L = \frac{\phi}{h}\left(y_{\zeta,t} - (1-h) n_{\zeta,t}^L\right) + \frac{\sigma_\zeta^H}{h}\left(y_{\zeta,t} - (1-h) c_{\zeta,t}^L\right). \tag{B.9}$$

To simplify we can obtain that:

$$\phi n_{\zeta,t}^L + (h\sigma_\zeta^L + (1-h)\sigma_\zeta^H) c_{\zeta,t}^L = (\phi + \sigma_\zeta^H) y_{\zeta,t}. \tag{B.10}$$

The budget constraint of the L household can be written as :

$$c^L_{\zeta,t} = w_{\zeta,t} + n^L_{\zeta,t} + \frac{\mu^d}{1-h} d_{\zeta,t}, \tag{B.11}$$

using $w_{\zeta,t} = -d_{\zeta,t} = \phi n^L_{\zeta,t} + \sigma^L_\zeta c^L_{\zeta,t}$, we can obtain that:

$$c^L_{\zeta,t} = (\phi n^L_{\zeta,t} + \sigma^L_\zeta c^L_{\zeta,t})(1 - \frac{\mu^d}{1-h}) + n^L_{\zeta,t}. \tag{B.12}$$

Substituting out $n^L_{\zeta,t}$, we can obtain that:

$$c^L_{\zeta,t} = ((\phi+\sigma^H_\zeta)y_{\zeta,t} - (h\sigma^L_\zeta + (1-h)\sigma^H_\zeta)c^L_{\zeta,t} + \sigma^L_\zeta c^L_{\zeta,t})(1 - \frac{\mu^d}{1-h}) + \frac{1}{\phi}(\phi+\sigma^H_\zeta)y_{\zeta,t} - (h\sigma^L_\zeta + (1-h)\sigma^H_\zeta)c^L_{\zeta,t}. \tag{B.13}$$

Let us first use the fact that $\sigma^H_e = \sigma^L_e = \sigma_e$ to simplify the expression as:

$$c^L_{e,t} = c^L_{e,t}\sigma_e \left(1 - \frac{\mu^d}{1-h}\right) + \chi^e \left(\frac{\phi+\sigma_e}{\phi} y_{e,t} - \frac{\sigma_e}{\phi} c^L_{e,t}\right). \tag{B.14}$$

Rearrange we can obtain that:

$$c^L_{e,t} = \chi^e y_{e,t}, \tag{B.15}$$

with

$$\chi^e = 1 + \phi(1 - \frac{\mu^d}{1-h}).$$

Now, we can do the same thing for the condition in recession such that:

$$c^L_{r,t} = c^L_{r,t}\sigma^L_r \left(1 - \frac{\mu^d}{1-h}\right) + \chi^e \left(\frac{\phi+\sigma^H_r}{\phi} y_{r,t} - \frac{(h\sigma^L_r + (1-h)\sigma^H_r)}{\phi} c^L_{r,t}\right). \tag{B.16}$$

Rearrange we can obtain that:

$$c^L_{r,t} = \chi^r y_{r,t}, \tag{B.17}$$

with

$$\chi^r = \frac{(\sigma^H_r + \phi)\chi^e}{\phi + \sigma^L_r + (1-h)\chi^e(\sigma^H_r - \sigma^L_r)}.$$

B.3 Deriving with fiscal policy

For fiscal policy, using the goods market condition and the resource constraints, we obtain that:

$$\phi n^L_{\zeta,t} + (h\sigma^L_\zeta + (1-h)\sigma^H_\zeta)c^L_{\zeta,t} = (\phi + \sigma^H_\zeta)y_{\zeta,t} - \sigma^H_\zeta g_{\zeta,t}. \tag{B.18}$$

And log linearization for the budget constraint of the low-income household:

$$c^L_{\zeta,t} = (\phi n^L_{\zeta,t} + \sigma^L_\zeta c^L_{\zeta,t})(1 - \frac{\mu^d}{1-h}) + n^L_{\zeta,t} - t^L_{\zeta,t}. \tag{B.19}$$

Combining two equations we obtain that:

$$c^L_{\zeta,t} = \chi^\zeta y_{\zeta,t} - \frac{\chi^\zeta \sigma^H_\zeta}{\phi + \sigma^H_\zeta} g_{\zeta,t} - \frac{\chi^\zeta}{\chi^e} \frac{\phi}{\phi + \sigma^H_\zeta} t^L_{\zeta,t}. \tag{B.20}$$

B.3.1 Deriving the IS curve

From the Euler equation of the high-income household, we have:

$$c^H_t = E_t c^H_{t+1} - \frac{1}{\sigma^H_\zeta}(i_t - E_t \pi_{t+1}). \tag{B.21}$$

Using the consumption respond of the low-income:

$$c^L_{\zeta,t} = \chi^\zeta y_{\zeta,t}. \tag{B.22}$$

With market clearing conditions $(1-h)c^L_t + hc^H_t$ to obtain the high-income consumption such that:

$$c^H_t = \frac{1 - (1-h)\chi^\zeta}{h} y_t. \tag{B.23}$$

Combining Eq. (B.23) and EqB.21, we can obtain the IS curve:

$$(\frac{1 - (1-h)\chi^\zeta}{h}) y_t = E_t (\frac{1 - (1-h)\chi^\zeta}{h}) y_{t+1} - \frac{1}{\sigma^H_\zeta}(i_t - E_t \pi_{t+1}). \tag{B.24}$$

Simplify to obtain that:

$$y_t = E_t y_{t+1} - \frac{\psi_{c,\zeta}}{\sigma^H_\zeta}(i_t - E_t \pi_{t+1}), \tag{B.25}$$

where

$$\psi_{c,\zeta} = \frac{h}{1 - (1-h)\chi^\zeta}.$$

As in the main text.

B.4 Deriving the IS curve with Fiscal Policy

Similar to the one without fiscal policy, using market clearing conditions and high-income consumption such that :

$$c_t^H = \frac{1-(1-h)\chi^\zeta}{h}(y_t - g_t) - \frac{1-h}{h}(\frac{\chi^\zeta \phi}{\phi + \sigma_\zeta^H} g_t - \frac{\phi}{\phi + \sigma_\zeta^H} t_t^L). \tag{B.26}$$

Combining with Euler's equation as in Eq. (B.21), similar to the previous process after rearrangement we can obtain that:

$$c_t = E_t c_{t+1} - \psi_{c,\zeta}(1-h)[\psi_{g,\zeta}(E_t g_{t+1} - g_t) - \psi_{\tau,\zeta}(E_t t_{t+1}^L - t_t^L)] - \frac{\psi_{c,\zeta}}{\sigma_\zeta^H}(i_t - E_t \pi_{t+1}), \tag{B.27}$$

where

$$\psi_{g,e} = \frac{\chi^e}{h}\left(\frac{\phi}{\phi + \sigma_e}\right),$$
$$\psi_{g,r} = \frac{\chi^r}{h}\left(\frac{\phi}{\phi + \sigma_r^H}\right),$$

$$\psi_{\tau,e} = \frac{1}{h}\left(\frac{\phi}{\phi + \sigma_e}\right),$$
$$\psi_{\tau,r} = \frac{1}{h}\frac{\chi^r}{\chi^e}\left(\frac{\phi}{\phi + \sigma_r^H}\right).$$

B.5 Aggregation of the NKPC Slope Across Types

The main text's Phillips-curve slopes use a population-weighted average of the type-specific curvatures. Because this aggregation step is not obvious when curvature is heterogeneous, we state it explicitly. Real marginal cost equals the wage under linear technology. The wage is common to both types, so it can be written as the population-weighted average of the two labor conditions:

$$w_{\zeta,t} = h\left(\phi n_{\zeta,t}^H + \sigma_\zeta^H c_{\zeta,t}^H\right) + (1-h)\left(\phi n_{\zeta,t}^L + \sigma_\zeta^L c_{\zeta,t}^L\right) = \phi n_{\zeta,t} + \bar{\sigma}_\zeta c_{\zeta,t} + h(1-h)\left(\sigma_\zeta^H - \sigma_\zeta^L\right)\left(c_{\zeta,t}^H - c_{\zeta,t}^L\right), \tag{B.28}$$

where $\bar{\sigma}_\zeta \equiv h\sigma_\zeta^H + (1-h)\sigma_\zeta^L$ and the last term is the covariance between curvature and consumption across types. Using $n_{\zeta,t} = y_{\zeta,t} = c_{\zeta,t}$ (no government spending) and the inequality identity $c_{\zeta,t}^H - c_{\zeta,t}^L = \frac{1-\chi^\zeta}{h} y_{\zeta,t}$, the exact marginal-cost loading on output is

$$w_{\zeta,t} = \left[\phi + \bar{\sigma}_\zeta + (1-h)\left(\sigma_\zeta^H - \sigma_\zeta^L\right)\left(1 - \chi^\zeta\right)\right] y_{\zeta,t}. \tag{B.29}$$

The resulting Phillips-curve slopes, used throughout the main text, are

$$\kappa_e = \frac{\kappa}{1+\phi}\big(\phi + 1 + (1-\alpha)\theta\big), \qquad \kappa_r = \frac{\kappa}{1+\phi}\left(\phi + 1 - (1-\alpha)\Big(h\tfrac{\theta}{\lambda^H} + (1-h)\tfrac{\theta}{\lambda^L}\Big)\right), \quad \text{(B.30)}$$

with government-spending pass-through $\kappa_{g,e} = \frac{\phi}{\phi+\sigma_e^H}$ and $\kappa_{g,r} = \frac{\phi}{\phi+\sigma_r^H}$, and $\kappa = (\frac{1}{\Theta} - 1)(1 - \beta\Theta)(\phi+1)$; since θ, λ^i, and h are all positive, $\kappa_e > \kappa_r$. In expansions the covariance term vanishes ($\sigma_e^H = \sigma_e^L$), and the slope reduces to κ_e exactly. In recessions κ_r, which is used throughout the analytics and the simulation, retains only $\phi + \bar{\sigma}_r$ (note that $\bar{\sigma}_r = 1-(1-\alpha)\theta\left(h/\lambda^H + (1-h)/\lambda^L\right)$ is the population-weighted expression in the main text) and drops the covariance term. At the baseline calibration the dropped term equals $(1-h)(\sigma_r^H - \sigma_r^L)(1-\chi^r) = 0.25 \times 0.218 \times (-0.642) = -0.035$, which would lower κ_r from 0.0493 to 0.0483 (about 2 percent) and flatten the recession Phillips curve slightly further. The omission is thus conservative for the paper's mechanism ($\kappa_r < \kappa_e$ holds a fortiori), and no result in the paper is sensitive to it.

B.6 Deriving the IS curve with idiosyncratic risk on changing type

Incorporating the fact that each agent perceives that their type may change with probability $1-s$, the Euler equation of the high-income household reads, in levels,

$$U_C(C_t^H, X_t) = \beta R_t E_t\left[\frac{P_t}{P_{t+1}}\Big(s\,U_C(C_{t+1}^H, X_{t+1}) + (1-s)\,U_C(C_{t+1}^L, X_{t+1})\Big)\right], \quad \text{(B.31)}$$

where the continuation marginal utilities are evaluated at the consumption *and the loss-aversion coefficient* of the type the household will have next period. Log-linearizing around the steady state with equal consumption across types, and using the fact that the log-deviation of marginal utility of a type-i household in regime ζ is $-\sigma_\zeta^i \hat{c}_t^i$, the continuation curvature is type- and regime-contingent:

$$c_t^H = sE_t c_{t+1}^H + (1-s)\frac{\sigma_{\zeta'}^L}{\sigma_\zeta^H} E_t c_{t+1}^L - \frac{1}{\sigma_\zeta^H}(i_t - E_t\pi_{t+1}), \quad \text{(B.32)}$$

where ζ' denotes the regime prevailing at $t+1$. Under the maintained assumption that the regime is absorbing within the response window ($\zeta' = \zeta$; see below), Eq. (B.32) specializes as follows. In expansions, $\sigma_e^L = \sigma_e^H$, so the curvature ratio is unity and the equation reduces *exactly* to the specification used in the simulation:

$$c_t^H = sE_t c_{t+1}^H + (1-s)E_t c_{t+1}^L - \frac{1}{\sigma_\zeta^H}(i_t - E_t\pi_{t+1}). \quad \text{(B.33)}$$

In recessions, the corrected equation weights the low-type continuation term by $\sigma_r^L/\sigma_r^H < 1$: because the low-type's effective curvature declines by more in the loss regime, the marginal

utility a switching household expects to have responds less than proportionally to its future consumption. Carrying the correction through the aggregation below replaces the forward-looking coefficient $\psi_{f,r}$ of Eq. (D.3) by

$$\tilde{\psi}_{f,r} = s + (1-s)\,\frac{\sigma_r^L}{\sigma_r^H}\,\frac{h\chi^r}{1-(1-h)\chi^r}, \tag{B.34}$$

which at the baseline calibration equals 1.021 against the uncorrected 1.044: the correction dampens the forward-compounding term in recessions by about two percentage points. The difference between the two specifications is bounded by $(1-s)\left(1-\sigma_r^L/\sigma_r^H\right) \approx 0.011$ per quarter of the continuation horizon, which is second order relative to the state-dependence effects the model is designed to capture; the simulated results reported in the main text are robust to this correction in the direction of *slightly weaker* recession amplification, so using the corrected equation would not overturn, and would marginally moderate, the quantitative asymmetries.

Two statements about regime handling complete the derivation. First, regimes in the simulation are dated by the position of consumption relative to the reference point ($\delta_t < 0$ is the loss regime) and are treated as absorbing within each impulse-response window, consistent with the calibrated shock sizes that do not switch the state (main text, Section 5); marginal utility is continuous at $\delta = 0$, so no boundary condition beyond the switch of coefficients is required. Second, the exact analytics of Section 4 avoid the issue entirely by setting types permanent ($s = s_L = 1$), for which Eq. (B.32) is exact regime by regime; the switching specification should be read as an approximation around that benchmark whose validity is governed by the bound above, and which we use only in the quantitative type-switching exercises of Section 5.

Again, using the consumption response conditions below:

$$\begin{cases} c^L_{\zeta,t} = \chi^\zeta(y_{\zeta,t} - g_{\zeta,t}) + h(\psi_{g,\zeta} g_t - \psi_{\tau,\zeta} t^L_{\zeta,t}) & \text{for } \zeta = [e,r], \\ c^H_{\zeta,t} = \frac{1-(1-h)\chi^\zeta}{h}(y_{\zeta,t} - g_{\zeta,t}) - (1-h)(\psi_{g,\zeta} g_t - \psi_{\tau,\zeta} t^L_{\zeta,t}) & \text{for } \zeta = [e,r], \end{cases} \tag{B.35}$$

Letting $c = y - g$ as the net output, we can obtain that:

$$\begin{aligned} c_t = {} & (1 + (\chi^\zeta - 1)(1-s)) E_t c_{t+1} \\ & - \psi_{c,\zeta}(1-h)\left[\psi_{g,\zeta}(E_t g_{t+1} - g_t) - \psi_{\tau,\zeta}(E_t t^L_{t+1} - t^L_t)\right] \\ & + \psi_{c,\zeta}(1-s)(\psi_{g,\zeta} E_t g_{t+1} - \psi_{\tau,\zeta} E_t t^L_{t+1}) \\ & - \frac{\psi_{c,\zeta}}{\sigma^H_\zeta}(i_t - E_t \pi_{t+1}), \end{aligned} \tag{B.36}$$

This equation converges to the TANK model as discussed in the main text when s=1. This condition reveals interesting dynamics under idiosyncratic risk: high-income agents may increase savings to buffer against the risk of downward mobility, while low-income agents might increase spending due to the prospect of upward mobility. The impact of these behaviors on future consumption depends on whether the consumption-output loading (χ) of low-income individuals is greater or less than 1, as explored in Bilbiie (2020) and Bilbiie (2025).

Furthermore, the potential for income mobility influences the efficacy of fiscal policy. An expansionary fiscal policy, which aims to boost aggregate demand, may have enhanced effects if low-income individuals, anticipating a transition to higher income brackets with a larger share of firm profits, increase their consumption. This amplifies the direct impact of government spending on current consumption levels.

C Additional Simulation Result

The simulation process is the same as in the main text, using the full model of Eq (B.36). We show the result for the extended model in which agents use habit formation as their consumption reference point. For the extended model, since the backward-looking behaviour after incorporating fiscal policy makes the derivation of the IS curve more complicated, we report the habit-reference results for monetary policy only, in Figure C.1. Indeed, after incorporating the habit reference point (i.e., $X_t = C_{t-1}^{\gamma}$), we obtain the IS curve below:

$$y_t = \vartheta_{1_\zeta}\Omega_\zeta y_{t+1} + \frac{\vartheta_{2_\zeta}(s(1-\chi^\zeta)+h\chi^\zeta)}{1-(1-h)\chi^e}y_{t-1} - \vartheta_{3_\zeta}(i-\pi_{t+1}), \tag{C.1}$$

where

$$\vartheta_{1_e} = \frac{1+(1-\alpha)\theta}{1+(1-\alpha)(1+\gamma)\theta}, \tag{C.2}$$

$$\vartheta_{2_e} = \frac{(1-\alpha)\gamma\theta}{1+(1-\alpha)(1+\gamma)\theta}, \tag{C.3}$$

$$\vartheta_{3_e} = \frac{h}{1-(1-h)\chi^e}\frac{1}{1+(1-\alpha)(1+\gamma)\theta}, \tag{C.4}$$

$$\Omega_e = (1+(\chi^e-1)(1-s)), \tag{C.5}$$

$$\vartheta_{1_r} = \frac{1-(1-\alpha)\theta\left(\frac{h}{\lambda^H}+\frac{1-h}{\lambda^L}\right)}{1-(1-\alpha)(1+\gamma)\theta\left(\frac{h}{\lambda^H}+\frac{1-h}{\lambda^L}\right)}, \tag{C.6}$$

$$\vartheta_{2_r} = \frac{(1-\alpha)\gamma\theta\left(\frac{h}{\lambda^H}+\frac{1-h}{\lambda^L}\right)}{1-(1-\alpha)(1+\gamma)\theta\left(\frac{h}{\lambda^H}+\frac{1-h}{\lambda^L}\right)}, \tag{C.7}$$

$$\vartheta_{3_r} = \frac{h}{1-(1-h)\chi^r}\frac{1}{1-(1-\alpha)(1+\gamma)\theta\left(\frac{h}{\lambda^H}+\frac{1-h}{\lambda^L}\right)}, \tag{C.8}$$

$$\Omega_r = (1+(\chi^r-1)(1-s)). \tag{C.9}$$

The NKPC will thus become:

$$\pi_t = E_t\beta\pi_{t+1} + \kappa_\zeta y_t + \kappa_{2,\zeta} y_{t-1} - \kappa a_t, \tag{C.10}$$

where

$$\kappa_{2,e} = (1-\alpha)\theta\gamma\kappa/(1+\phi), \tag{C.11}$$

$$\kappa_{2,r} = \frac{(1-\alpha)\theta\gamma\kappa\left(\frac{h}{\lambda^H}+\frac{1-h}{\lambda^L}\right)}{1+\phi}. \tag{C.12}$$

C.1 Model-Data Comparison

Figure C.2 compares the simulated and empirical responses of earnings inequality, as discussed in Section 5.2.

D Counter-cyclical Income Risk

In this section, we analyze the case in which the probability of switching type is endogenized into the decision-making of the household. Two distinct models appear in this appendix, and we state at each point which one a given equation describes: (i) the *loss-aversion model with switching* (our full model), in which the regime-dependent curvatures σ^i_ζ operate and the switching probability is state-dependent; and (ii) the *income-risk rival*, in which loss aversion is shut down entirely ($\alpha = 1$, so by Eq. (17) the effective curvature is $\sigma^i_\zeta = \sigma = 1$ in both states with no gain-loss kink, and $\chi^\zeta = \chi^e$) and *all* state dependence comes from an exogenously counter-cyclical switching probability $(1 - s_\zeta)$. The horse race of the online Appendix D.1 compares (i) against (ii). Intuitively, if high-income households know they have some probability of switching type and becoming low-income households, with a smaller dividend share and less income, this is

Figure C.1: Comparing results under AR(2) specification with extended model: Consumption Inequality



Note: The figure contrasts the impacts of policy measures during recessions and expansions, captured as $\beta^r - \beta^e$, in response to a 1% contractionary monetary policy shock on consumption inequality (left panel) and earning inequality (right panel). The simulated data (blue line) are derived from the extended-model simulation over 5,000 impulse response periods, using the local projection method with AR(2) and MA(20) specified in Eq. (1), alongside the empirical results of Section 2.2 (red dotted line) with confidence intervals (green vertical lines).

reflected in their intertemporal choice between today's and tomorrow's consumption. In model (i), loss aversion additionally makes high-income households more sensitive to a given switching risk in recessions; in model (ii), the counter-cyclicality of the risk is imposed directly from the evidence of Guvenen et al. (2014).

To see more directly why this generates asymmetric, counter-cyclical income risk, we follow the same decomposition as Bilbiie (2025), first looking at the conditional mean of the total income of the high-income household:

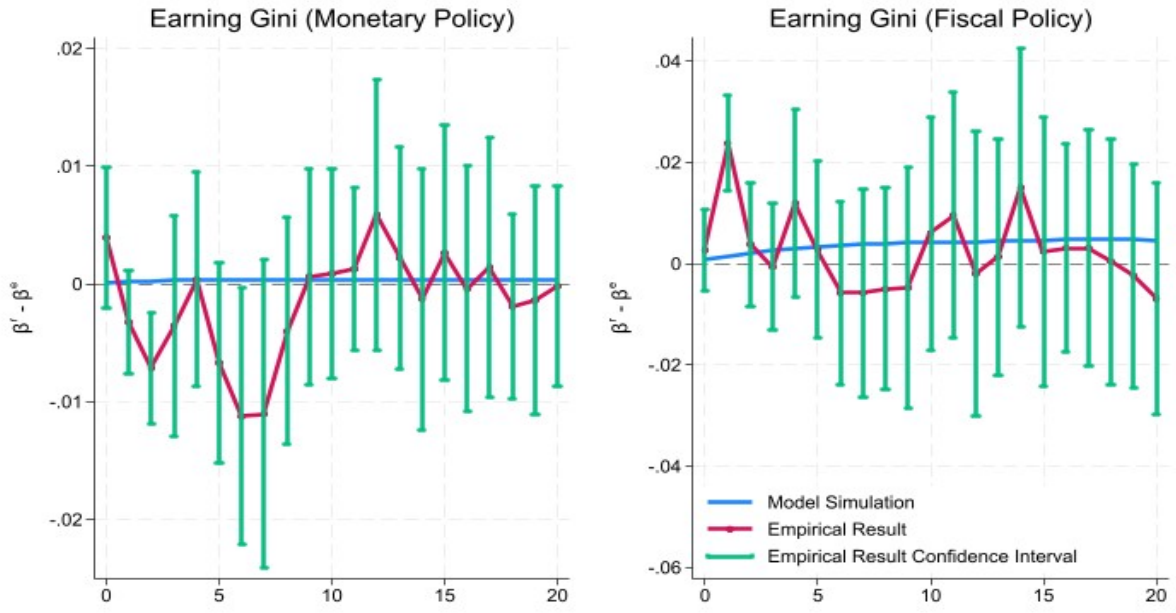
$$E_t\left(\ln Y_{t+1}^H \mid \ln Y_t^H\right) = s \ln Y_{t+1}^H + (1-s) \ln Y_{t+1}^L \tag{D.1}$$

Using the two-state version of the Rouwenhorst method we obtain:

$$\text{var}\left(\ln Y_{t+1}^H \mid \ln Y_t^H\right) = s(1-s)\left(\ln \frac{Y_{t+1}^H}{Y_{t+1}^L}\right)^2 = s(1-s)\,\gamma_{t+1}^2 \tag{D.2}$$

where $\gamma_{t+1} = \ln \frac{Y_{t+1}^H}{Y_{t+1}^L} = \ln \frac{C_{t+1}^H}{C_{t+1}^L} = c_{\zeta,t+1}^H - c_{\zeta,t+1}^L$ is the consumption-inequality gap introduced in Section 4.3. Since we have already shown that the consumption inequality generated by loss

Figure C.2: Comparing Empirical and Simulated Results for Earning Inequality



Note: The figure illustrates the differential impacts of policy measures during recessions and expansions, quantified as $\beta^r - \beta^e$, in response to a 1% contractionary monetary policy shock (left panel) and a 1% expansionary fiscal policy shock (right panel) on earning inequality, defined in the model as the sum of wage and dividend income (empirically, both labor earnings and total before-tax income show no robust asymmetry). This analysis employs the local projection method with AR(2) and MA(20) specified in Eq. (1) and simulates the model over 5,000 impulse response periods, represented by the blue line. The empirical results of Section 2.2 are shown by the red dotted line with confidence intervals (green vertical lines).

aversion is counter-cyclical and larger during recessions, the same holds here: our model also explains why income risk is larger for high-income households in recessions than in expansions.

Aggregate Demand with Idiosyncratic Risk To understand how income risk affects aggregate demand, we derive a generalized IS curve incorporating both state-dependent loss aversion and switching probabilities. Defining net output as $c = y - g$, we obtain:

$$
\begin{aligned}
c_t = \psi_{f,\zeta} E_t c_{t+1} & \\
& - \psi_{c,\zeta}(1-h)\left[\psi_{g,\zeta}(E_t g_{t+1} - g_t) - \psi_{\tau,\zeta}(E_t t^L_{t+1} - t^L_t)\right] \\
& + \psi_{c,\zeta}(1-s)(\psi_{g,\zeta} E_t g_{t+1} - \psi_{\tau,\zeta} E_t t^L_{t+1}) \\
& - \frac{\psi_{c,\zeta}}{\sigma^H_\zeta}(i_t - E_t \pi_{t+1})
\end{aligned} \tag{D.3}
$$

with:

$$
\psi_{f,\zeta} = (1 + (\chi^\zeta - 1)\frac{(1-s)}{1-(1-h)\chi^\zeta}), \quad \psi_{c,\zeta} = \frac{h}{1-(1-h)\chi^\zeta} \tag{D.4}
$$

The coefficient $\psi_{f,\zeta}$ captures how expected future consumption affects current spending decisions through two channels. First, the term χ^ζ reflects the state-dependent consumption response to income changes. Second, the switching probability $(1-s)$ determines how strongly households internalize the risk of income changes.

The calibration of the switching probabilities $(1-s_e, 1-s_r)$ (targeting the average quarterly transition rate of Kaplan et al. (2014) and the cyclical left-tail risk and skewness evidence of Guvenen et al. (2014), yielding $1-s_e = 4.5\%$ and $1-s_r = 6.8\%$) is described in the online Appendix D.1, together with the resulting comparison in Figure D.1.[26] This appendix provides the underlying derivations and the analytical decomposition behind the quantitative statements made there.

Asymmetric Effect and Income Risk Now, since we use a tractable New Keynesian model from Bilbiie (2025), this allows us to demonstrate analytically why the counter-cyclical income risk fails to match the pattern of loss-aversion quantitatively.

For analytical purposes, we focus on the case when there is no fiscal policy and decompose the dynamic impact of the output and interest as in Bilbiie (2025) and Pfäuti and Seyrich (2022) to obtain that:

$$
c_{\zeta,t} = \frac{1-\beta(1-(1-h)\chi^\zeta)}{1-\beta\psi_{f,\zeta}\rho(1-(1-h)\chi^\zeta)} y_t - \frac{h\beta}{\sigma^H_\zeta(1-\beta\psi_{f,\zeta}\rho(1-(1-h)\chi^\zeta))} r_t, \tag{D.5}
$$

where ρ is the persistence of the monetary policy. The first term ($\frac{1-\beta(1-(1-h)\chi^\zeta)}{1-\beta\psi_{f,\zeta}\rho(1-(1-h)\chi^\zeta)} = \Gamma_{GE,\zeta}$) is the indirect effect of the monetary policy on consumption via output. The second term

[26] In the aggregation of the average transition rate, $1-\bar{s} = \omega_e(1-s_e) + \omega_r(1-s_r) = 0.054$, the weights $\omega_e = 0.85$ and $\omega_r = 0.15$ are the ergodic probabilities of the expansion and recession states, calibrated to the historical U.S. business cycle patterns documented by the NBER: the economy spends approximately 85% of the time in expansion and 15% in recession during the post-WWII period.

($\frac{h\beta}{\sigma_\zeta^H(1-\beta\psi_{f,\zeta}\rho(1-(1-h)\chi^\zeta))} = \Gamma_{D,\zeta}$) is the direct monetary policy effect on consumption. Applying the state-dependence inequality we previously discussed, we can obtain that:

$$c_{\zeta,t}^H - c_{\zeta,t}^L = \frac{1-\chi^\zeta}{h} c_{\zeta,t} = \frac{1-\chi^\zeta}{h}(\Gamma_{GE,\zeta} y_t - \Gamma_{D,\zeta} r_t). \tag{D.6}$$

Now, the important question is: without loss aversion, how much of the asymmetric monetary policy effect on consumption inequality can be explained? To address this question, we compare the ratios of the direct and indirect effects across the business cycle. Since the income-risk term is $\psi_{f,\zeta} = (1 + (\chi^\zeta - 1)\frac{(1-s)}{1-(1-h)\chi^\zeta})$, we can address this question by assuming the low-income consumption loading is fixed across the business cycle while the probability of switching type is larger during recessions than expansions, consistent with the empirical observation of Guvenen et al. (2014) that tail risk is larger during recessions. We compare the case in which there is no switching risk and the asymmetry is driven purely by loss aversion with the case in which there is no loss aversion and the asymmetry is driven by the switching probability.

In the calibration of the online Appendix D.1, the benchmark expansion and recession switching risks are 4.5% and 6.8% respectively. However, to generate the same magnitude of asymmetric effects observed in our general equilibrium model without loss aversion, the recession-state switching probability would need to rise from its calibrated 6.8% (4.5% in expansions) to 47.2%. For the direct effect alone, the probability would need to increase to 9.2% in the left-tail income risk during recessions.

One might ask why such high uncertainty is necessary to replicate the asymmetry explained by loss aversion alone. First, the loss-aversion mechanism impacts not only the intertemporal trade-off between present and future consumption but also the consumption-labor choice itself. In contrast, while idiosyncratic income risk may deter high-income individuals from consuming as much as lower-income individuals during a recession, it does not alter labor market decisions in the same way. This difference arises because precautionary savings, driven by future income uncertainty, contrasts with loss aversion, which is triggered by the current economic environment. Second, loss aversion affects all households by reducing their consumption-labor trade-off, but lower-income, hand-to-mouth households face binding consumption constraints, adjusting primarily in response to aggregate demand shifts driven by high-income households. This results in asymmetry largely driven by the consumption adjustments of high-income households.

In sum, our analysis demonstrates that adding idiosyncratic income risk explains (1) why income risk escalates in recessions, (2) why output volatility intensifies in downturns, and (3)

Figure D.1: Comparing Loss Aversion and Idiosyncratic Risk



Note: The figure compares the effects of loss aversion (blue line) and idiosyncratic income risk (green line) on the differential impact of policy measures during recessions and expansions, quantified as $\beta^r - \beta^e$, in response to a 1% contractionary monetary policy shock, for consumption inequality (left panel) and earning inequality (right panel). This analysis employs the local projection method with AR(2) and MA(20) specified in Eq. (1) and simulates each model over 5,000 impulse response periods. The empirical results of Section 2.2 are shown by the red dotted line with confidence intervals (vertical bars).

why loss aversion remains essential to fully capture the asymmetric response to policy shocks across the business cycle. Figure D.1 plots the resulting comparison, discussed in the online Appendix D.1.

D.1 A Horse Race with Counter-Cyclical Income Risk

To give the income-risk alternative its best shot, we endogenize the switching probability and calibrate its cyclical variation to Guvenen et al. (2014): an average quarterly transition rate of $1-\bar{s}=0.054$ (Kaplan et al., 2014), a 50 percent increase in left-tail risk in recessions ($\frac{1-s_r}{1-s_e}=1.5$), and increased negative skewness ($\frac{(1-s_r)/s_r}{(1-s_e)/s_e}=2.0$), yielding $1-s_e=4.5\%$ and $1-s_r=6.8\%$ (with s_L adjusting to hold $1-h=0.25$). This average switching rate exceeds the baseline model's, which works in the income-risk model's favor. The online Appendix D presents the generalized IS curve and full derivations.

Figure D.1 shows the result side by side. Both models deliver a near-flat earning-inequality response (right panel), so the earnings margin does not discriminate between them; the dis-

criminating statistic is the *magnitude* of the consumption-inequality asymmetry each mechanism generates. On that statistic the income-risk model moves in the right direction but falls well short of the empirical asymmetry that the loss-aversion model matches at its baseline saver/hand-to-mouth gradient. Loss aversion operates through the effective curvature of preferences given *current* consumption relative to reference, so its full force arrives exactly in recessions without requiring any asymmetry in relative earnings—the configuration in the data.

The two mechanisms are complements: adding switching risk to the loss-aversion model changes the fit little, but switching risk without loss aversion misses the magnitude.

This comparison does not displace income risk and liquidity constraints as the workhorse channels; it shows that, within the linear-quadratic approximation of Section 4.3, part of the effectiveness of recession policy support may reflect households' greater responsiveness once consumption falls below reference (a lower effective curvature, not a rise in loss aversion).

E Welfare-Based Optimal Policy

E.1 Derivation of the Welfare Loss Function

We derive the central bank's welfare loss function from a second-order approximation to household utility.

E.1.1 Social Welfare Function

The social welfare function aggregates utility across both household types:

$$\mathcal{W} = E_0 \sum_{t=0}^{\infty} \beta^t \left[h\mathcal{U}_t^H + (1-h)\mathcal{U}_t^L \right], \tag{E.1}$$

where the period utility for household type $i \in \{H, L\}$ is:

$$\mathcal{U}_t^i = U(C_t^i, X_t^i) - \frac{(N_t^i)^{1+\phi}}{1+\phi}, \tag{E.2}$$

with $U(C_t^i, X_t^i)$ being the loss aversion utility function (Eqs. (2)–(3)) and ϕ the inverse Frisch elasticity.

From the loss aversion specification, the marginal utility of consumption for household i in state ζ is:

$$U_C^i = \frac{1}{C_t^i} \left[\alpha + (1-\alpha) G_C(C_t^i, X_t^i) \right]. \tag{E.3}$$

Linearizing around steady state with reference point $X_t^i = \bar{C}$, the log-deviation of marginal utility is:

$$\hat{U}_C^i \approx -\sigma_\zeta^i \hat{c}_t^i, \tag{E.4}$$

consistent with the Euler equation (21), in which the real rate enters scaled by $(\sigma_\zeta^H)^{-1}$, where σ_ζ^i is the state-dependent effective curvature parameter given in Eq. (17):

$$\sigma_e^H = 1 + (1-\alpha)\theta, \quad \sigma_e^L = 1 + (1-\alpha)\theta \tag{E.5}$$

$$\sigma_r^H = 1 - (1-\alpha)\frac{\theta}{\lambda^H}, \quad \sigma_r^L = 1 - (1-\alpha)\frac{\theta}{\lambda^L}. \tag{E.6}$$

The marginal disutility of labor is:

$$U_N^i = -(N_t^i)^\phi \approx -\bar{N}^\phi(1 + \phi\hat{n}_t^i). \tag{E.7}$$

Taking a second-order Taylor expansion of the welfare function around an efficient flexible-price steady state and using the equilibrium conditions, we obtain (following Woodford, 2003):

$$\mathcal{W} \approx \bar{\mathcal{W}} - \frac{1}{2}E_0\sum_{t=0}^{\infty}\beta^t\Bigg[\text{Inflation losses} + \text{Output gap losses} \tag{E.8}$$

$$+ \text{Consumption dispersion losses}\Bigg] + \text{t.i.p.}. \tag{E.9}$$

Inflation losses: From Calvo price stickiness with parameter Θ and elasticity of substitution ϵ:

$$\text{Inflation losses} = \frac{\epsilon}{\xi}\pi_t^2. \tag{E.10}$$

where $\xi = \frac{(1-\Theta)(1-\beta\Theta)}{\Theta}$.

Output gap losses: From labor disutility and consumption volatility (per period, before normalization):

$$\text{Output gap losses} = \left(\phi + (\sigma_\zeta^H)^{-1}\right) y_t^2. \tag{E.11}$$

The term ϕ captures labor disutility (from $\sum_i U_N^i \hat{n}_t^i$), while $(\sigma_\zeta^H)^{-1}$ captures the curvature of marginal utility for high-income households who determine the aggregate Euler equation.

Consumption dispersion losses: The welfare loss from unequal consumption across types carries the effective curvature itself (the second-order term of utility is $\frac{1}{2}\sigma_\zeta^i(\hat{c}_t^i)^2$ after normalization by the marginal utility of aggregate consumption):

$$\text{Dispersion losses} = h\,\sigma_\zeta^H(\hat{c}_t^H)^2 + (1-h)\,\sigma_\zeta^L(\hat{c}_t^L)^2. \tag{E.12}$$

Using the decomposition $h(\hat{c}_t^H)^2 + (1-h)(\hat{c}_t^L)^2 = \hat{c}_t^2 + h(1-h)(\hat{c}_t^H - \hat{c}_t^L)^2$, the aggregate part is absorbed into the output-gap term, and (setting $\sigma_\zeta^L \approx \sigma_\zeta^H$, which is exact in expansions and an approximation in recessions) the cross-sectional part yields the dispersion weight used below.

Using the equilibrium conditions from Eq. (14) and market clearing:

$$\hat{c}_t^L = \chi^\zeta \hat{y}_t, \quad \hat{c}_t^H = \frac{1-(1-h)\chi^\zeta}{h}\hat{y}_t. \tag{E.13}$$

After algebraic manipulation, the dispersion losses can be written in terms of consumption inequality $\gamma_t = c_t^H - c_t^L = \frac{1-\chi^\zeta}{h} y_t$:

$$\text{Dispersion losses} = \sigma_\zeta^H h(1-h)\,\gamma_t^2. \tag{E.14}$$

Normalizing so that inflation carries a unit coefficient (the convention behind the weights used in the main text, derived explicitly in the reconciliation paragraph below) we obtain:

$$\mathcal{L}^\zeta = \frac{1}{2} E_0 \sum_{t=0}^{\infty} \beta^t \left[\pi_t^2 + \alpha_y^\zeta y_t^2 + \alpha_\gamma^\zeta \gamma_t^2\right], \tag{E.15}$$

where:

$$\alpha_y^\zeta = \frac{\phi + (\sigma_\zeta^H)^{-1}}{\epsilon(1-\beta\Theta)} \tag{E.16}$$

$$\alpha_\gamma^\zeta = \frac{\sigma_\zeta^H h(1-h)}{\epsilon(1-\beta\Theta)} = \sigma_\zeta^H h(1-h)\alpha_y^\zeta. \tag{E.17}$$

Substituting the equilibrium relationship $\gamma_t = \frac{1-\chi^\zeta}{h} y_t$ into equation (E.15):

$$\mathcal{L}^\zeta = \frac{1}{2} E_0 \sum_{t=0}^{\infty} \beta^t \left[\pi_t^2 + \alpha_y^\zeta y_t^2 + \alpha_\gamma^\zeta \left(\frac{1-\chi^\zeta}{h}\right)^2 y_t^2\right] \tag{E.18}$$

$$= \frac{1}{2} E_0 \sum_{t=0}^{\infty} \beta^t \left[\pi_t^2 + \Lambda^\zeta y_t^2\right], \tag{E.19}$$

where:

$$\Lambda^\zeta = \alpha_y^\zeta \left[1 + \sigma_\zeta^H h(1-h)\left(\frac{1-\chi^\zeta}{h}\right)^2\right]. \tag{E.20}$$

E.1.2 First-Order Conditions

Set up the Lagrangian:

$$\mathcal{L}_t = \frac{1}{2}\left[\pi_t^2 + \Lambda^\zeta y_t^2\right] + \Xi_t\left[\pi_t - \beta E_t[\pi_{t+1}] - \kappa_\zeta y_t - u_t\right], \tag{E.21}$$

where Ξ_t is the Lagrange multiplier on the Phillips curve (we reserve λ for the loss-aversion coefficient).

First-order conditions:

$$\frac{\partial \mathcal{L}_t}{\partial \pi_t}: \quad \pi_t + \Xi_t = 0. \tag{E.22}$$

$$\frac{\partial \mathcal{L}_t}{\partial y_t}: \quad \Lambda^\zeta y_t - \Xi_t \kappa_\zeta = 0. \tag{E.23}$$

Eliminating $\Xi_t = -\pi_t$ yields the targeting rule:

$$\pi_t = -\frac{\Lambda^\zeta}{\kappa_\zeta} y_t. \tag{E.24}$$

Substituting this into the IS curve with $g_t = 0$:

$$y_t = E_t[y_{t+1}] - \frac{\psi_{c,\zeta}}{\sigma_\zeta^H}(i_t - E_t[\pi_{t+1}]). \tag{E.25}$$

Using $y_t = -\frac{\kappa_\zeta}{\Lambda^\zeta}\pi_t$ and $E_t[y_{t+1}] = -\frac{\kappa_\zeta}{\Lambda^\zeta}E_t[\pi_{t+1}]$:

$$-\frac{\kappa_\zeta}{\Lambda^\zeta}\pi_t = -\frac{\kappa_\zeta}{\Lambda^\zeta}E_t[\pi_{t+1}] - \frac{\psi_{c,\zeta}}{\sigma_\zeta^H}(i_t - E_t[\pi_{t+1}]). \tag{E.26}$$

Multiplying by $-\frac{\Lambda^\zeta}{\kappa_\zeta}$:

$$\pi_t = E_t[\pi_{t+1}] + \frac{\psi_{c,\zeta}\Lambda^\zeta}{\sigma_\zeta^H \kappa_\zeta}(i_t - E_t[\pi_{t+1}]). \tag{E.27}$$

Rearranging:

$$\pi_t - E_t[\pi_{t+1}] = \frac{\psi_{c,\zeta}\Lambda^\zeta}{\sigma_\zeta^H \kappa_\zeta}(i_t - E_t[\pi_{t+1}]), \tag{E.28}$$

$$(\pi_t - E_t[\pi_{t+1}])\frac{\sigma_\zeta^H \kappa_\zeta}{\psi_{c,\zeta}\Lambda^\zeta} = i_t - E_t[\pi_{t+1}], \tag{E.29}$$

$$i_t = E_t[\pi_{t+1}] + \frac{\sigma_\zeta^H \kappa_\zeta}{\psi_{c,\zeta}\Lambda^\zeta}(\pi_t - E_t[\pi_{t+1}]), \tag{E.30}$$

$$i_t = E_t[\pi_{t+1}]\left(1 - \frac{\sigma_\zeta^H \kappa_\zeta}{\psi_{c,\zeta}\Lambda^\zeta}\right) + \frac{\sigma_\zeta^H \kappa_\zeta}{\psi_{c,\zeta}\Lambda^\zeta}\pi_t. \tag{E.31}$$

For a forward-looking rule (responding to expected inflation), set the coefficient on current inflation to zero by choosing the appropriate timing. The implementable rule is:

$$i_t = \phi^\zeta E_t[\pi_{t+1}], \tag{E.32}$$

where the policy coefficient is:

$$\phi^\zeta = 1 + \frac{\sigma_\zeta^H \kappa_\zeta}{\psi_{c,\zeta}\Lambda^\zeta}. \tag{E.33}$$

E.1.3 State-Dependence and the Taylor Principle

The policy coefficient ϕ^ζ is state-dependent through all its components:

$$\phi^\zeta = 1 + \frac{\sigma_\zeta^H \kappa_\zeta}{\psi_{c,\zeta}\Lambda^\zeta} = 1 + \frac{\sigma_\zeta^H \kappa_\zeta}{\frac{h}{1-(1-h)\chi^\zeta} \cdot \alpha_y^\zeta \left[1 + \sigma_\zeta^H h(1-h)\left(\frac{1-\chi^\zeta}{h}\right)^2\right]}. \tag{E.34}$$

Simplifying:

$$\phi^\zeta = 1 + \frac{\sigma_\zeta^H \kappa_\zeta [1-(1-h)\chi^\zeta]}{h\alpha_y^\zeta \left[1 + \sigma_\zeta^H h(1-h)\left(\frac{1-\chi^\zeta}{h}\right)^2\right]}. \tag{E.35}$$