# The Curious Incidence of Monetary Policy Across the Income Distribution*

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**Abstract**

We use high-frequency German administrative data to study the effects of monetary policy across the earnings distribution. Earnings growth at the bottom is roughly three times as responsive as at the median, driven by more countercyclical job separation rates among lower-earnings workers, while job-finding rates respond uniformly. In a HANK model, this unequal incidence amplifies consumption responses to monetary policy by a factor of two, through both heightened precautionary savings and larger consumption drops upon job loss, relative to an economy with homogeneous hours adjustment. Both unemployment risk itself and its concentration among high-MPC workers contribute to this amplification.

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# 1 Introduction

How do monetary policy interventions affect individuals' earnings and employment prospects across the income distribution? Does unequal incidence of monetary policy across the distribution amplify or dampen the response of aggregate consumption to changes in interest rates? The burgeoning heterogeneous-agent New Keynesian (HANK) literature has identified labor income as an important channel through which household heterogeneity impacts the transmission of monetary policy (inter alia, Auclert, 2019; Bilbiie, 2025; Hagedorn et al., 2019; Kaplan et al., 2018). Answers to the foregoing questions are therefore key for understanding the transmission of monetary policy to the aggregate economy. However, there is little direct empirical evidence from large advanced economies on these transmission channels, and similarly little guidance from structural macroeconomic frameworks as to their quantitative relevance.

In this paper, we first empirically study the heterogeneous effects of monetary policy surprises on labor earnings across the income distribution. Our findings show that monetary policy has significantly larger effects on the earnings of low-income workers. This is mainly because their job-loss risk responds more strongly to interest rate changes than that of workers with higher incomes. This unequal incidence significantly reduces income inequality during monetary expansions. It has long-lasting effects on the employment rates of poor workers, which remain elevated even years after the initial shock. Second, we use a structural model to show that the heterogeneous incidence of monetary policy on employment risk across the income distribution strongly amplifies its effect on aggregate demand relative to economies with hours adjustment or homogeneous unemployment risk, with the combined effect of these features roughly doubling the cumulative consumption response.

For our empirical analysis, we use a long panel of detailed German administrative data, containing individual labor market biographies including earnings. Labor market status is observed at a daily frequency. The high-frequency nature of our data allows us to estimate responses of earnings and labor market transition probabilities to monetary policy shocks and high-frequency changes in aggregate earnings. This sets our paper apart from the literature that empirically investigates the heterogeneous effects of business cycles on individual income risk using administrative data. Our dataset allows us to understand the importance of changes in employment status for earnings changes. The previous literature has speculated that the larger sensitivity of earnings at the bottom of the distribution was due to non-employment risk; our paper, to our knowledge, is the first to show that this is indeed the case.

We identify monetary policy surprises using high-frequency changes in Overnight Indexed

Swap (OIS) rates for the Eurozone.[1] We then use the identified shock series for estimating the impact of shocks on labor earnings using local projections à la Jordà (2005). Monetary policy affects labor earnings most at the bottom of the permanent-income distribution. In response to an expansionary monetary policy surprise, earnings growth rises about three times as much in the bottom quintile as it does at the median of the distribution. The differential growth is accounted for by a substantially stronger fall in separation rates into non-employment at the bottom of the distribution. In contrast, job-finding rates, while pro-cyclical, rise homogeneously across the distribution. Similarly, the earnings growth of workers who remain employed increases, but with mostly uniform effects across the income distribution.

These heterogeneous earnings responses across the distribution give rise to strong redistributional forces. An unexpected interest rate cut leads the Gini coefficient of labor earnings to fall significantly. In addition, monetary policy has significant effects on medium-run employment prospects: individuals who become unemployed in the month of a monetary policy expansion find jobs significantly faster, have significantly higher earnings and remain employed significantly longer.

To understand the implications of our empirical findings for the aggregate economy, we embed cyclical and heterogeneous unemployment risk into an otherwise standard heterogeneous-agent New Keynesian (HANK) model. Unemployment risk responds endogenously to aggregate fluctuations, in line with the empirical regularities we document: separation rates rise and job-finding rates fall in response to contractionary shocks, with both responses concentrated among low-income workers.

We use the model to isolate two sources of amplification through counterfactual analysis. First, we compare our benchmark economy, in which labor adjustment occurs through hiring and firing, with one in which the same aggregate labor input adjustment occurs through hours worked by the employed, holding the distribution of risk fixed. This comparison reveals that cyclical unemployment risk amplifies the consumption response to monetary policy through two channels: a precautionary-savings channel, as households reduce consumption when the probability of job loss rises, and a realized-income channel, as a rise in unemployment reduces consumption by more than an equivalent reduction in earnings from hours worked. Our decomposition suggests these channels are roughly equally important.

Second, we compare our benchmark economy, in which unemployment risk is concentrated among low-income workers, with one in which all workers face identical employment transition probabilities. This comparison reveals that heterogeneous incidence further amplifies the

[1] See e.g. Jarociński and Karadi (2020); Nakamura and Steinsson (2018), and Almgren et al. (2022) for discussion of high-frequency identification of monetary policy shocks in the U.S. and Eurozone, respectively.

consumption response, because monetary policy disproportionately affects high-MPC workers who account for the bulk of precautionary savings.

Combining these comparisons, unemployment risk and its heterogeneous incidence together raise the consumption response to monetary policy by roughly 100 percent relative to an economy with homogeneous hours adjustment—that is, one featuring neither unemployment risk nor its concentration among the income poor.

## Relation to the literature

Our paper contributes to four broad strands of the literature. First, it contributes to the empirical literature that studies the cyclical nature of income risk, beginning with Storesletten et al. (2004). Our paper fits into the more recent focus of this literature that uses high-quality administrative data (Guvenen et al. (2014, 2017); Patterson (2023) for the US, and for other countries (Halvorsen et al. (2024) (Norway), Hoffmann and Malacrino (2019) (Italy), De Nardi et al. (2019), (Netherlands and US)). We contribute to this literature by providing evidence from another large advanced economy (Germany), and by showing that the cyclical incidence of risk is the same for business cycle shocks as well as aggregate earnings movements induced by identified monetary policy shocks.

Second, our paper adds to the fast-growing literature studying how monetary policy affects households and individuals. An early important contribution is (Coibion et al., 2012), which studies the impact of monetary policy shocks using data from the Consumer Expenditure Survey in the US. Three contemporaneous papers—Holm et al. (2021), Andersen et al. (2023), and Amberg et al. (2022) investigate this question for the cases of Norway, Denmark, and Sweden, respectively. They all find similar patterns: earnings are more strongly affected at the low end of the income distribution than they are around the median. At the top of the distribution, results are mixed. Our research documents the same pattern in Germany. The novelty of our contribution is decomposing the mechanisms underlying the result. We show that conditional on employment there is a small homogeneous response of earnings. We then show that the U-shaped pattern is driven by labor market transitions, and in particular the separation margin. Two other papers written subsequent to ours also investigate labor-market implications to monetary policy. Coglianese et al. (2022) studies a single monetary policy event in Sweden after the Global Financial Crisis and looks at the implications for unemployment. Cantore et al. (2022) study labor supply effects of monetary policy in the U.K. and U.S. economies. Finally, a closely related paper, Moser et al. (2021) also uses administrative data to study how negative interest rates following the European Debt Crisis in 2014 impacted credit supply to firms and the employment and wage prospects of their workers. We see this paper as complementary to ours, in that it exploits a particular event study to better understand the mechanisms driving the labor-market transitions that we document.

Third, our paper contributes to the literature studying how heterogeneous incidence of labor income risk shapes the transmission of aggregate shocks. Auclert (2019) shows that the redistribution channel of monetary policy depends on how income gains and losses are distributed across households with different marginal propensities to consume. Patterson (2023) quantifies the "matching multiplier": the amplification that arises because recessions disproportionately reduce earnings for high-MPC workers. Similarly, Hagedorn et al. (2017, 2022) develop a general methodology to quantify the demand amplification arising from redistribution, across labor income, capital income, taxes, and transfers, induced by aggregate shocks relative to a complete-markets benchmark. Those papers study heterogeneous incidence through the intensive margin: workers remain employed but experience differential earnings changes. Our contribution is to show that the extensive margin—unemployment risk—is the primary driver of heterogeneous incidence in response to monetary policy, and that this distinction matters quantitatively.

Fourth, our paper contributes to the literature that studies how unemployment risk can amplify aggregate fluctuations. The pioneering work of Krusell and Smith (1998) studied idiosyncratic and aggregate unemployment risk in a real model with self-insurance. More recent contributions have focused on the importance of nominal rigidities, including Acharya and Dogra (2020); Challe (2020); Den Haan et al. (2018); Gornemann et al. (2016); Graves (2025); Ravn and Sterk (2017, 2021). We contribute by quantifying how the heterogeneous incidence of unemployment risk, its concentration among low-income, high-MPC workers, further amplifies the precautionary-savings channel emphasized in this literature.

The rest of the paper is organized as follows. Section 2 describes the administrative data and the structure of income and employment transitions in our sample. Section 3 explains how we identify monetary policy surprises and estimate their effects using local projections. Section 4 presents our main empirical results on the heterogeneous incidence of monetary policy across the earnings distribution. Section 5 develops the structural model and describe the results of our counterfactuals. Section 6 concludes.

# 2 Data

We use administrative social security data on a two-percent sample of all labor-market histories in Germany from the Sample of Integrated Employment Biographies (provided by the Research Data Center, FDZ).[2] This dataset contains about 1.7 million individuals but excludes civil servants and self-employed individuals. For our analysis, we utilize data on the years between 1995 and 2013. Each observation in the original dataset is a labor-market spell (Ganzer et al., 2017).[3] For our purposes, we convert these spells into monthly employment histories for each individual. Each such observation includes an individual's employment status and their average daily labor earnings, which we aggregate to the monthly level. Earnings are deflated using the Harmonized Index for Consumer Prices for Germany.[4] For about ten percent of individuals in our sample, earnings are top coded; we exclude these observations. All non-employed workers are coded to have zero income.

Because we are interested in the effect of monetary policy on labor earnings and employment status, we focus on individuals with a high degree of attachment to the labor market. In particular, we restrict our sample to employed individuals liable to social security without special characteristics, (thus excluding, for example, trainees and marginal part-time workers) and the unemployed, defined as individuals who received unemployment benefits (ALG I) at the beginning of their current non-employment episode. The usual definition of unemployment includes a provision about active job search, about which we have no information in our data. Hence, our unemployment definition is likely narrower that the one used to compute official unemployment statistics.

We study the differences in the earnings responses to monetary policy across the income distribution. We classify individuals according to their permanent income, which we see as a key summary measure of welfare differences, but also because previous work has found strong heterogeneity along this dimension (Guvenen et al., 2017). Our preferred proxy for permanent income is average earnings over the five years preceding the month for which we calculate the earnings change, as in Guvenen et al. (2017).[5] Using this measure, in every such month, we sort individuals into quantiles, conditional on gender and five-year age brackets. We restrict the sample to workers who have at least one earnings observation in the above-mentioned five-year period, in order to avoid bunching at zero. Furthermore, we exclude individuals whose earnings exceed the maximum amount liable for social security contributions, because their incomes are top-coded at that limit. For these workers, we cannot compute reliable

[2]We rely on the factually anonymous version of the Sample of Integrated Labour Market Biographies (SIAB-Regionalfile) – Version 7514. Research Data Centre (FDZ) of the Federal Employment Agency (BA) at the Institute for Employment Research (IAB). Data access was provided via a Scientific Use File supplied by the FDZ of the BA at the IAB.

[3]Employment relationships longer than 12 months are split into multiple spells. We drop spells that are shorter than 1 month. Potentially missing spells are imputed according to Drews et al. (2007).

[4]Obtained from Eurostat, series prc_hicp_midx.

[5]Our estimation sample comprises the period between 2000M1 to 2012M12. However, we make use of data from 1995 in order to compute our backward-looking permanent income measure, but only consider monetary policy surprises from 2000M1 to 2013M12.

earnings growth rates. After applying these sampling restrictions, we are left with close to 800,000 individuals across our sample period and around 300,000 individuals in each month.

To understand how key observables evolve along our permanent income distribution (henceforth simply the "income distribution"), Table 1 reports descriptive statistics across deciles for the month of January 2010.[6]

Table 1: Averages within deciles of permanent income, first quarter 2010

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | mean | mean | mean | mean | mean | mean | mean | mean | mean | mean |
| HS or Technical College | 0.04 | 0.04 | 0.03 | 0.03 | 0.03 | 0.03 | 0.04 | 0.04 | 0.06 | 0.08 |
| Vocational training | 0.72 | 0.75 | 0.78 | 0.81 | 0.82 | 0.83 | 0.82 | 0.83 | 0.81 | 0.74 |
| University | 0.09 | 0.07 | 0.06 | 0.05 | 0.05 | 0.05 | 0.06 | 0.06 | 0.08 | 0.15 |
| Daily wage, imputed | 1342.08 | 1574.44 | 1801.57 | 2013.29 | 2220.12 | 2443.86 | 2703.84 | 3004.74 | 3412.02 | 4253.54 |
| Monthly earnings | 1357.58 | 1583.56 | 1807.73 | 2019.15 | 2223.41 | 2447.28 | 2706.19 | 3006.75 | 3414.18 | 4257.66 |
| Employed | 0.81 | 0.87 | 0.92 | 0.95 | 0.97 | 0.98 | 0.98 | 0.99 | 0.99 | 0.98 |
| Job finding | 0.36 | 0.50 | 0.51 | 0.55 | 0.56 | 0.56 | 0.57 | 0.54 | 0.56 | 0.53 |
| Job loss | 0.07 | 0.06 | 0.04 | 0.03 | 0.02 | 0.02 | 0.02 | 0.01 | 0.01 | 0.02 |
| Observations | 29258 | 29230 | 29239 | 29242 | 29237 | 29243 | 29241 | 29240 | 29240 | 29233 |

**Note:** The table shows values of different variables averaged within deciles of the permanent income distribution in January 2010. Deciles are computed conditionally on five-year age brackets and gender. We impute education following the imputation procedure in Fitzenberger et al. (2005). Monthly earnings are in current Euros, all others are fractions. Job-finding and job-loss refer to $U$ to $E$ and $E$ to $U$ transitions over twelve months, respectively. The deciles are computed conditional on age and gender. These variables are thus not reported.

In our dataset, education is measured by a categorical variable. More than 70 percent of all individuals in our sample indicate vocational training as their highest qualification, and education levels are very similar across the first 8 deciles. In the last two deciles, the share of university educated individuals rises but never exceeds 20%. The gradient of nominal earnings across the distribution is substantial, with average earnings in the top decile about four times higher than in the first. In 2010, the social security contribution limit was about 5400 euros per month, beyond which we drop individuals. This explains the seemingly low value of average earnings in the top decile. Employment rates are high in this sample of highly-attached individuals. They average 76 percent in the bottom decile, and rise steeply across the bottom half of the distribution to flatten out around 98 percent above the median. Job-finding rates (defined as 12-month transitions of the unemployed into employment) are between 50 and 60 percent in all deciles but the first, where they are substantially lower (about a third). Job-loss probabilities (similarly defined) fall monotonically, from seven to two percent, across the distribution.

[6]Note that, with some abuse of language but hopefully no room for confusion, we call deciles both the 9 points of the distribution as well as the 10 groups they define (we proceed similarly for other quantiles). The quantiles are computed conditional on age and gender. These variables are thus not reported in Table 1.

# 3 Estimation strategy

## 3.1 Identifying monetary policy surprises

We focus on the period between January 2000 and December 2013, when European monetary policy was conducted by the ECB.[7] Since the German economy accounts for roughly one-quarter of Euro-area GDP it is likely that the ECB's monetary policy was heavily influenced by German economic performance. Hence, when estimating the impact of interest rate changes on the German economy, endogeneity is an important concern.

To identify monetary policy surprises our approach follows Almgren et al. (2022), who rely on high-frequency changes in Overnight Indexed Swap (OIS) rates. We use these changes to instrument for unexpected changes in the ECB's policy rate, which we denote as $\Delta i_t$. Every six weeks, on Thursdays, the ECB governing council meets to decide on monetary policy actions. At 13:45 CET, a press release is posted, which concisely summarizes the decisions taken by the governing council. Subsequently, at 14:30 CET, the president of the ECB holds a press conference, first motivating the decisions taken in an introductory statement and later taking questions from the audience. Our instrument, $Z_t$, equals the change in 3-month EONIA OIS rates in response to each of these two events in a narrow time window around them. If this measure is large, in absolute terms, we conclude that the decisions taken by the ECB Governing Council were not expected by financial markets and vice versa. The identifying assumption underlying the approach is that no other news are released during the above-mentioned short time windows which have an impact on the effectiveness of monetary policy.[8]

Our main empirical specification to estimate the effects of monetary policy surprises on economic variables is the following regression:

$$x_{t+h} - x_{t-1} = \alpha_h + \beta_h \Delta i_t + \gamma_h X_{t-1} + \varepsilon_{t,h} \tag{1}$$

where $x_{t-1}$ represents the value of the economic variable in question one period before the monetary policy surprise, and $x_{t+h}$ represents its value $h$ periods after the shock. We condition this growth rate on $x_{t-1}$, as opposed to $x_t$, because it is conceivable that monetary policy has contemporaneous effects on $x_t$, which would invalidate all growth rates going forward. The vector $X_{t-1}$ represents a set of control variables consisting of three lags of (i) the instrument $Z_t$, (ii) aggregate earnings and (iii) $\Delta i_t$, as well as calendar month dummies to control for

[7] The high-frequency identification approach outlined here cannot be implemented for earlier time periods, as the Bundesbank did not relay its policy decision on a precisely planned schedule on the announcement day.

[8] For more information, see Almgren et al. (2022).

seasonality. As the policy rate, we use the Euro Overnight Index Average (EONIA).[9]

### 3.2 Aggregate effects of monetary policy

Before moving to individual incomes, we investigate the effect of monetary policy shocks on the aggregate economy in Germany. To this end, we estimate the regression in Equation (1), replacing $x$ with (i) the logarithm of the consumer price index for Germany (OECD, 2025), (ii) the logarithm of industrial production (Eurostat, 2025a), (iii) the official German unemployment rate (Eurostat, 2025b), and (iv) the real interest rate, computed as the change in the logarithm of the German price index between months $t+1$ and $t$ subtracted from the EONIA rate in period $t$; we also augment the vector of control variables $X_t$ with three lags of the left-hand side variable.

Figure 1 shows the impulse responses to an contractionary monetary policy surprise, equal to one standard deviation of the policy rate (following Gertler and Karadi, 2015), estimated using Equation (1). The horizontal axis measures the time after the monetary policy shock in months, the vertical axis measures the percentage point change in the variable in question. The top left panel indicates that the inflation rate does not significantly react to the surprise in either direction. The response of GDP is reported in the top right panel. According to the textbook theory of monetary policy, production should contract following a monetary tightening. The graph indicates that this is the case. The unemployment rate (bottom left panel) increases slowly but significantly so. The real interest rate increases after the monetary policy shock, but then returns to zero after about 1.5 years. Most of the subsequent estimates are insignificant.

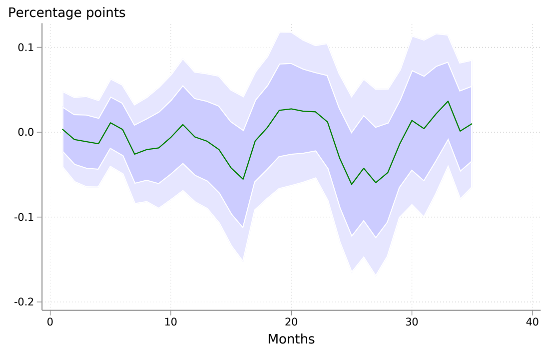
In addition to standard macro aggregates, Figure 15 in the appendix plots the change in aggregate earnings (i.e., average earnings across all individuals) and the employment share for our full sample. The left panel shows that the response of earnings to a contractionary monetary policy surprise of one-standard-deviation builds up gradually, reaching a point estimate of about 0.5 percentage points after two years. This reduction in average earnings is accompanied by a fall in the share of workers who are employed.

## 4 Monetary-policy effects across the income distribution

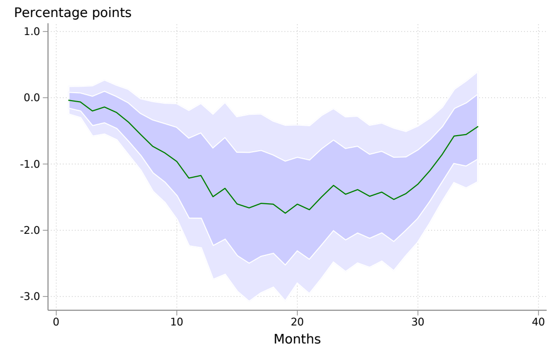
In this section, we estimate the effect of monetary policy surprises on earnings growth and its drivers, in particular movements in and out of employment.

[9] The EONIA was discontinued in 2022. We obtain a time series for it from the replication package of Almgren et al. (2022).

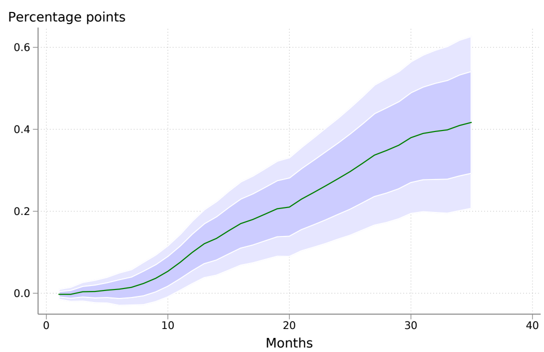
Figure 1: Aggregate responses to monetary policy surprises



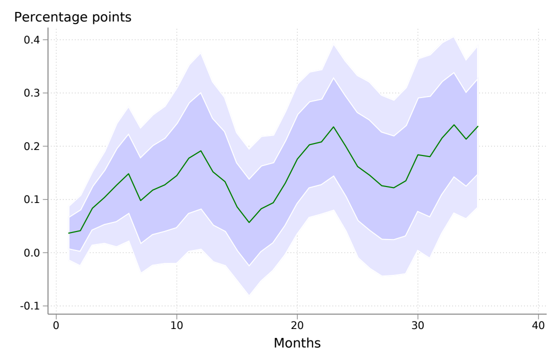

(a) Inflation rate




(b) Industrial production




(c) Unemployment rate




(d) Real interest rate

**Note:** The figure shows the impulse responses of aggregate variables to a surprise contractionary shock of one standard deviation, estimated using the LPIV outlined in Equation (1). The *Top Left Panel* shows the change in the inflation rate, calculated as the change in the logarithm of the HICP for Germany. The *Top Right Panel* shows the percentage change in industrial production, calculated as the log difference between $t$ and $t + h$, and the *Bottom Left Panel* shows the change in the unemployment rate. The *Bottom Right Panel* shows the change in the real interest rate, calculated as the inflation rate subtracted from the policy rate. The sample period is from 2000 until 2013. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC standard errors.

## 4.1 Earnings growth and employment rates

We start by estimating the effect of monetary policy surprises on earnings growth rates, separately for individuals in different quantiles of the permanent income distribution. As described in Section 2, we sort individuals into quantiles based on their permanent income in period $t - 1$. We split the distribution into 20 quantiles, or ventiles, conditioning on age and gender. For each of these quantile groups, we first compute average earnings as

$$\overline{earn}^q_{t+h} = \frac{1}{N^q} \sum_{i=1}^{N_q} earn_{i,t+h} \quad \forall i \in q \text{ at } t-1$$

where $earn_{i,t+h}$ represents the labor earnings of an individual in month $t+h$ who was sorted into quantile $q$ in month $t-1$.

Next, in Equation (1), we set $x_{t+h} = \log(\overline{earn}_{t+h})$ and, for each quantile, estimate

$$\Delta \log(\overline{earn}^{q}_{t+h}) = \alpha_h + \beta^q_h \Delta i_t + \theta X_t + \epsilon^q_{t+h} \tag{2}$$

where $\Delta x_{t+h} = x_{t+h} - x_{t-1}$. The coefficient $\beta^q_h$ captures the effect of a change in interest rates $\Delta i_t$, in period $t$, (instrumented by $Z_t$, as described in Section 3, following Stock and Watson (2018)) on earnings growth in quantile $q$ between periods $t-1$ and $t+h$ .

We scale the size of the exogenous interest rate change, $\Delta i$, such that it causes a 1 percentage point increase in the growth rate of *aggregate* earnings, twelve months after the shock.[10] This allows us to compare the change in earnings growth rates across quantiles associated with an *unconditional* one-percent change in aggregate earnings (as in Guvenen et al. (2017)) to that of a *conditional* change in aggregate earnings of equal size, caused by a monetary policy innovation:

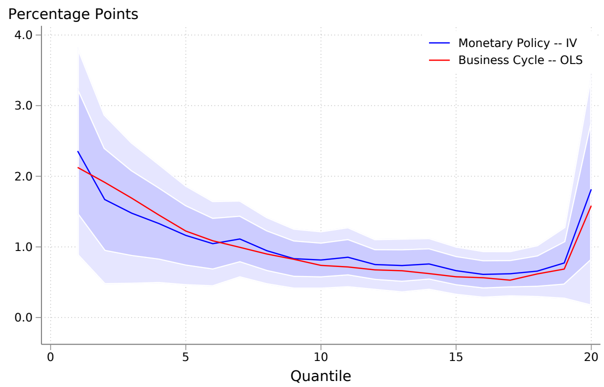
$$\Delta \log(\overline{earn}^{q}_{t+h}) = \alpha_{Y,h} + \beta^q_{Y,h} \Delta \log(\overline{earn}_{t+h}) + \theta X_t + \epsilon^q_{Y,t+h}. \tag{3}$$

Here, the coefficient $\beta^q_{Y,h}$ represents the change in the quantile-specific earnings average in response to an unconditional change in the overall earnings average.

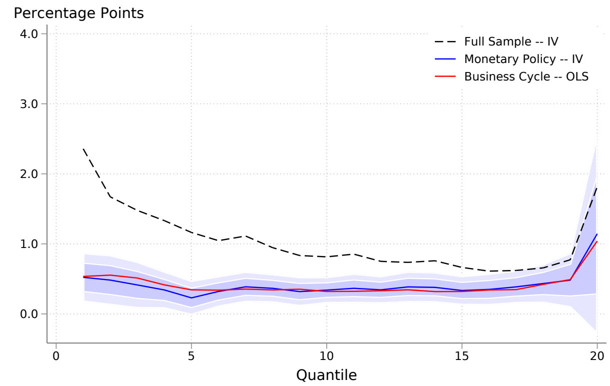
The blue line in the left panel of Figure 2 reports the quantile-specific earnings changes between months $t-1$ and $t+12$, induced by an exogenous interest rate change which raises aggregate earnings by one percentage point over the same period. Recall that, since the maximum length of an employment spell in our dataset is twelve months, earnings growth between $t-1$ and $t+12$ is always computed using earnings observations drawn from two different employment spells.

[10] For reference, the left panel of Figure 15 plots the response of aggregate earnings in our sample to an exogenous *one-standard-deviation* rise in the policy rate. Aggregate earnings fall by roughly 0.4 percentage points. Hence, to induce a 1 percentage point rise in aggregate earnings, the policy rate must fall by more than twice as much.

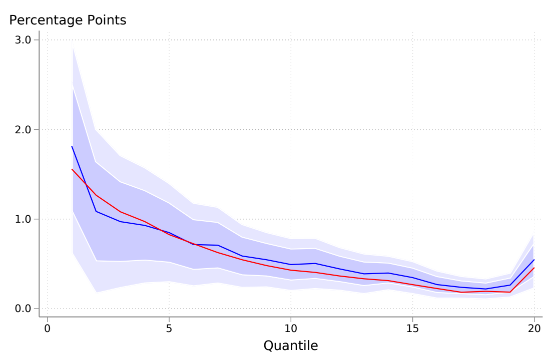
Figure 2: Regression coefficients $\beta_{12}^{q}$ across the income distribution




(a) Earnings growth, full sample




(b) Earnings growth, individuals employed in $t-1$ and $t+12$




(c) Change in share of employed workers

**Note:** The *Top Left Panel* plots the coefficients $\beta_{12}^{q}$ in Equation (2) (scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings) and $\beta_{Y,12}^{q}$ in Equation (3), separately for individuals who shared the same ventile of the permanent income distribution in period $t-1$. The *Top Right Panel* compares the coefficients $\beta_{12}^{q}$ for the full sample across ventiles (gray dashed line) to $\beta_{12}^{q,E}$ and $\beta_{Y,12}^{q,E}$, estimated on a smaller sample of individuals who are employed both in period $t-1$ and $t+12$ (the blue and red lines, respectively). The *Bottom Panel* plots the coefficients $\beta_{12}^{q}$ in Equation (2), where the left-hand side is the average employment rate by quantile. Ventiles are constructed based on average earnings during the five years prior to $t-1$, conditional on gender and five-year age brackets. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC robust standard errors. The sample period is 2000-2013.

Earnings changes in response to expansionary monetary policy exhibit a pronounced U-shape across the permanent income distribution. In particular, the earnings of the poorest individuals, in the bottom ventile, respond almost three times as much as earnings at the median. Moving up the income distribution, this response declines strongly in magnitude, to about two-thirds of the median effect in ventiles 15 to 19.[11] Earnings of the income-rich, in the

[11]In Appendix B.7, we dicuss which coefficients are significantly different from each other.

top ventile, respond more, about twice as strong as median earnings.[12] Both Andersen et al. (2023) and Amberg et al. (2022) find similar patterns for Denmark and Sweden, respectively; Cantore et al. (2022) documents a similar pattern in the US using data from the CPS.[13]

The red line in the left panel of Figure 2 depicts the point estimates $\beta^q_{Y,12}$, summarising the comovement of individual and aggregate earnings growth rates without conditioning on monetary policy surprises. As documented in Guvenen et al. (2017) for the US economy, this comovement also has a U-shaped relationship with the level of individual permanent incomes, very similar to that of earnings changes due to monetary policy (although with a somewhat less pronounced increase in the extreme ventiles).

The estimates of $\beta^q_{12}$, depicted in the left panel of Figure 2, conflate the effect monetary policy has on labor earnings with the effect it has on employment probabilities, as they are based on the changes in average labor earnings of all individuals in a given quantile (including the unemployed who have zero labor earnings). However, because average earnings across quantiles, $\overline{earn}^q_t$, equal the product of the average labor earnings of the employed, $\overline{earn}^{q,E}_t$ times the employment rate, we can compute the following decomposition

$$\log(\overline{earn}^q_{t+h}) = \log\left(\overline{earn}^{q,E}_{t+h}\right) + \log\left(\frac{N^{q,E}_{t+h}}{N^q_{t+h}}\right) \tag{4}$$

where $\overline{earn}^{q,E}_{t+h}$ represents the average earnings of employed individuals in month $t+h$, who were sorted into quantile $q$ in period $t-1$. In the second expression, $N^{q,E}_{t+h}$ represents the number of employed individuals in period $t+h$ who were sorted into quantile $q$ in period $t-1$. Thus, Equation (4) implies that changes in average labor earnings across quantiles are the sum of two separate effects: the changes in the labor earnings of the employed (which we denote the *intensive*-margin effect), and changes in the employment rate (*extensive*-margin effect).

To isolate the heterogeneity in the intensive-margin effect, we substitute the change in average earnings of the employed, $\overline{earn}^{q,E}_{t+h}$, in place of its full-sample counterpart $\overline{earn}^q_{t+h}$ in Equation (2). The resulting coefficients, which we refer to as $\beta^{q,E}_{12}$, are displayed in the blue line in the right panel of Figure 2. As before, we scale the point estimates such that the initial exogenous interest rate change $\Delta i$ causes aggregate earnings growth to rise by one percentage point over the subsequent twelve months.

[12] Our results are robust to the inclusion of unemployment benefits (see Appendix 4.5), to alternative definitions of permanent income (see Appendix B.1) and to different measures of monetary policy surprises (see Appendix B.2)

[13] In Andersen et al. (2023), the U-shape in earnings responses is less pronounced. We conjecture that this is due to the fact that in their approach, individuals are sorted into an income distribution based on incomes including transfers. Hence, their distribution includes individuals that are only partially attached to the labor force, especially towards the bottom. Our approach excludes those individuals.

Earnings of the employed appear to be much less affected by monetary policy surprises than earnings in the full sample. The estimates are less heterogeneous across the distribution and substantially smaller in magnitude. In response to an exogenous change to the policy rate, the earnings growth of the employed rises by about 0.7 percentage points in the first quantile. The point estimate of this effect declines somewhat across the first five quantiles, but is essentially flat between ventiles 9 and 19, before rising substantially (but not significantly) in the top ventile. The difference between the estimates of $\beta_{12}^{q}$ (dashed black line) and $\beta_{12}^{q,E}$ is most pronounced in the bottom ventile, where the extensive margin of employment accounts for two thirds of monetary policy's effect on average labor earnings. This role of the extensive margin declines across the income distribution, to about a quarter of the overall effect.

In order to document the effect of a monetary expansion on employment by quantiles, the second term in the decomposition (4), we estimate Equation 2, substituting the log change in average employment for the log change in average earnings on the left-hand side. The bottom panel in Figure 17 shows the results of this exercise. At the bottom of our income distribution, employment rises by about two percentage points, while towards the top, employment is almost unchanged. Again, these results imply that monetary policy affects the low end of the income distribution more strongly than the top, although our estimates are imprecisely estimated in the exercise at hand.

In the next subsection, we further investigate the drivers of the extensive-margin effect of monetary policy by studying its effect on transition frequencies across different labor market states.

## 4.2 Labor market transitions

We observe each individual in our sample either as employed or as unemployed. Let $s_{i,1}$ be an individual's labor market status in period $t-1$ and $s_{i,2}$ be the labor market status of the same individual in some future period $t+h$. Then, there are four different transitions between $s_{i,1} \in \{E, U\}$ and $s_{i,2} \in \{E, U\}$. In addition, we also identify a subset of "switchers" who are observed as employed in both periods, but with different employers ($s_2 = switch$).

For each quantile along the permanent-income distribution, we aggregate the individual transitions into transition probabilities:

$$TR_{t+h}^{q,s_1,s_2} = \frac{1}{N^{q,s_1}} \sum_{i \in q,s_1} \mathbb{I}_{s_1,s_2}.$$

According to this definition, $TR_{t+h}^{q,s_1,s_2}$ is the fraction of all individuals who are sorted into quantile $q$ in period $t-1$ and observed in state $s_1$ at $t-1$ ($N^{1,s_1}$), who have transitioned to state $s_2$ by period $t + h$.

Similarly to Equation (2), we then estimate the following regression separately for each quantile-subsample:

$$TR^{q,s_1,s_2}_{t+h} = \alpha + \gamma^{q,s_1,s_2}_h \Delta i_t + \theta X_t + \epsilon^{q,s_1,s_2}_{t+h} \tag{5}$$

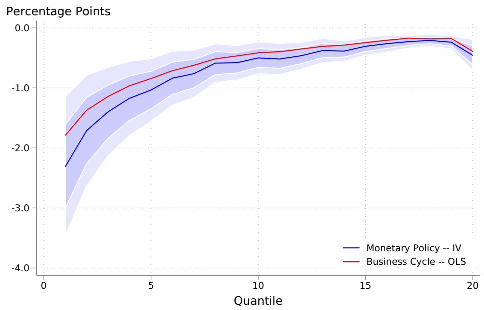
where the coefficient $\gamma^{q,s_1,s_2}_h$ measures the percentage point change in the share of individuals in state $s_1$ that make a particular labor market transition in response to a monetary policy surprise, for a given quantile $q$. Again, the vector $X_t$ contains calendar-month dummies and three lagged values of $\Delta i_t$, aggregate earnings, and $Z_t$.

The blue line in the top left panel of Figure 3 shows the point estimates for $\gamma^{q,E,U}_{12}$ (again scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings), summarising the effect of a monetary policy surprise on transitions from employment to unemployment. As with earnings, we document strong heterogeneity in the incidence of monetary policy surprises along the income distribution. For the poorest individuals in the sample, the interest rate change significantly decreases the probability of moving to unemployment by on average two percentage points. Moving up the income distribution this effect declines monotonically to less than 0.5 percentage points. The top ventile is again affected somewhat more strongly.

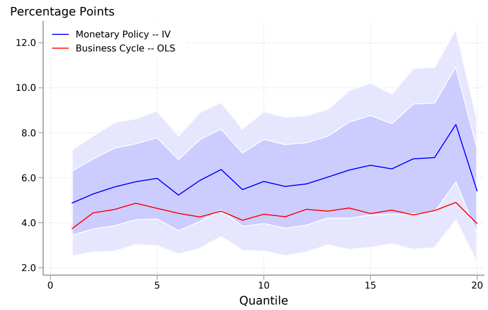
Analogous to section 4.1, we can compare the estimates conditional on monetary policy with those of unconditional comovement between transition probabilities and aggregate earnings changes.[14] The resulting coefficients are displayed as the red line in the top left panel of Figure 3. Interestingly, the reduction in transitions into unemployment is somewhat more pronounced for the expansionary monetary policy shock than for an unconditional increase in aggregate earnings. The difference is largest at the low end of the permanent income distribution.

[14]The regression is of the same form as Equation 5, with $\Delta i$ substituted for with changes in aggregate income $\Delta Y$. We label the resulting coefficient $\gamma^{q,s_1,s_2}_{Y,h}$.

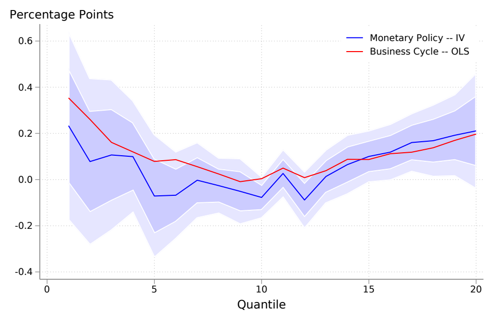
Figure 3: Regression coefficients $\gamma_{12}^{q}$ across the income distribution




(a) E to U




(b) U to E




(c) Switch

**Note:** The *Top Left Panel* plots the coefficients $\gamma_{12}^{q,E,U}$ in Equation (5) (scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings, blue line) and $\gamma_{Y,12}^{q,E,U}$ (red line), from a version of Equation (5) which quantifies unconditional comovement (see text). Both quantify the change in transition probabilities for the employed in $t-1$ to unemployment in period $t+12$ (E to U). The *Top Right Panel* plots the scaled coefficients $\gamma_{12}^{q,U,E}$, and $\gamma_{Y,12}^{q,U,E}$, for the share of unemployed transiting to employment (U to E). The *Bottom Panel* plots the scaled coefficient $\gamma_{Y,12}^{q,switch}$ and $\gamma_{Y,12}^{q,switch}$ for the share of the employed who change employment relation. Ventiles are constructed based on average earnings during the five years prior to $t-1$, conditional on gender and five-year age brackets. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC standard errors. The sample period is 2000-2013.

The top right panel of Figure 3 shows the scaled point estimates for $\gamma_{12}^{q,U,E}$, summarising the effect of an expansionary monetary policy surprise on the probability of unemployed individuals transitioning to employment. This effect is on average more than 5 percentage points. Contrary to the stronger effect on the likelihood of E-to-U transitions, U-to-E transitions respond slightly less to monetary policy at the bottom of the distribution. In particular, while monetary policy shocks affect the transition probabilities of the income-poor similarly to average fluctuations (as summarised by their comovement with average earnings, in the red line), a gap between the two opens up along the income distribution.

The results in the top panels of Figure 3 thus show that the substantially stronger extensive-margin effect of monetary policy on employment shares of the poor is largely accounted for by their decline in employment-to-unemployment transitions. The bottom panel of Figure 3 further investigates the source of this heterogeneity. It shows the scaled point estimates for $\beta^{switch}_{TR,12}$, summarising the effect of monetary policy surprises on the frequency of transitions between two different employment relationships. An expansionary monetary policy surprise makes job-switching more likely in the bottom quartile, but has small or imprecisely measured effects in the rest of the distribution. A similar pattern holds for the effect on job-switching of unconditional fluctuations in average earnings.

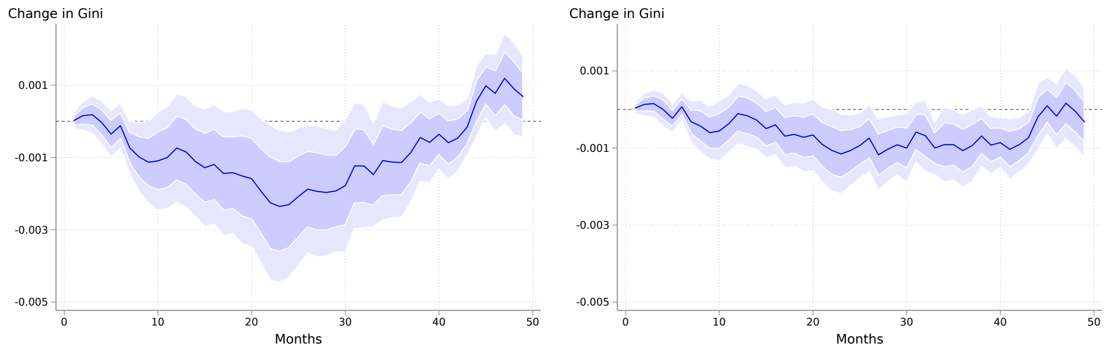
## 4.3 Inequality

The previous results beg the question how inequality in labor earnings develops in response to changes in monetary policy. To investigate this, we substitute values of the aggregate monthly Gini coefficient, $gini_{t+h}$, for $x$ in Equation (1). Importantly, we include unemployed individuals in our calculations, with their labor earnings set to zero, as above.

The left panel in Figure 4 plots the change in the Gini coefficient in response to an expansionary monetary policy shock over time.[15] Inequality falls for two years after the shock, then reverts back. Throughout our sample period, the average value of the Gini coefficient is close to 0.3, implying that monetary policy has economically significant effects on this measure, decreasing it by close to one percent at the trough of the impulse response function in Figure 4.

[15] As before, the monetary policy surprise is scaled to cause aggregate earnings to rise by one percentage point over twelve months.

Figure 4: Gini coefficient Impulse Response




(a) Change in Gini

(b) Change in Gini, including benefits

**Note:** The *Left Panel* shows the change in the Gini coefficient of labor earnings (including zeros), *gini*, in response to an expansionary monetary policy surprise, consistent with a one-percent increase in aggregate earnings, over time. The *Right Panel* shows the change in the Gini coefficient of labor earnings, including unemployment benefit receipts, $gini^{UI}$, in response to an expansionary monetary policy surprise, consistent with a one-percent increase in aggregate earnings, over time. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC standard errors. The sample period is 2000-2013.

Because our dataset also includes some information about unemployment benefit receipts, we can calculate the Gini coefficient taking these benefits into account. Analogous to before, we substitute $gini^{UI}_{t+h}$ into Equation (1) and compute the impulse response of this statistic to an expansionary monetary policy surprise. The implied change in inequality, plotted in the right panel of Figure 4, is substantially smaller, compared to the case when the unemployed's earnings are set to zero. Although $gini^{UI}$ decreases significantly, after about two years, the change is economically small: less than 0.5 percent relative to its average value of 0.25. The unemployment benefit system, therefore, appears to attenuate the effect that monetary policy has on inequality.

## 4.4 Earnings and labor market prospects after unemployment

Figure 3 shows that much of the effect of monetary policy on average earnings, and most of its heterogeneous incidence, is due to the response of labor market transitions between employment and unemployment. However, the previous section only investigates a short, 12-month window of labor market transitions. Because the costs of unemployment are strongly affected by its duration and effect on future earnings, this section discusses the longer-run effects of monetary policy shocks on re-employment probabilities and earnings after unemployment. For this, we compare two groups of individuals: workers who become unemployed in the period of the monetary surprise ($t$) and those who retain their jobs in

$t$. We then investigate how average employment and earnings of the second group evolve relative to the first, and how monetary policy affects the difference. For $h = -6, -5, ..., 36$ we run the following regression for three terciles of our permanent income distribution:

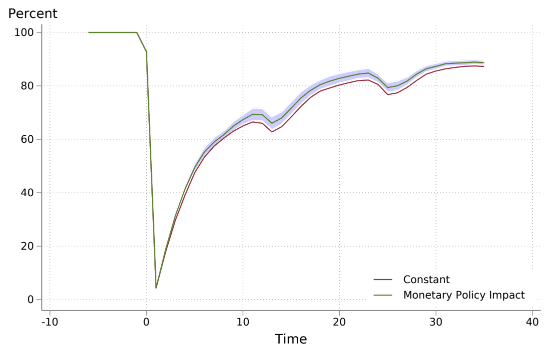
$$x_{t+h} = \alpha_{x,h} + \gamma_{x,h}\Delta i_t + \theta_{x,h} X_t + \epsilon_{x,t}. \tag{6}$$

where, $x \in \{earn, emp\}$ corresponds to the ratio of (i) the average employment rate $\overline{emp}_{i,t}$ or (ii) average monthly earnings ($\overline{earn}_{i,t}$) between individuals who become unemployed in period $t$ and those who do not. Again, $\Delta i_t$ represents the interest rate change in period $t$, instrumented using $Z_t$ as before, and $X_t$ contains calendar-month dummies and three lags of the interest rate change, the instrument and aggregate earnings. In Equation (6), $\alpha_{x,h}$ captures the average earnings or employment ratio $h$ months after an unemployment shock in the absence of monetary policy surprises, between the two groups. In turn, $\gamma_{x,h}$ quantifies the impact of monetary policy on these variables. The regressions are similar in spirit to that in Davis and Von Wachter (2011), who also investigate earnings paths of the unemployed relative to those who remain employed. We focus on individuals who become unemployed after being employed for at least 6 months. Because this substantially reduces the sample size, we report results for terciles, rather than ventiles, of the permanent income distribution.

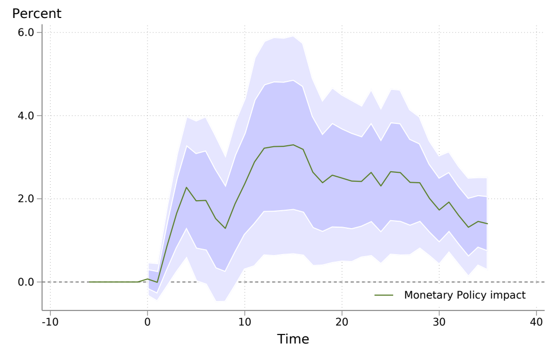
The top panels of Figure 5 show the results of the exercise for employment probabilities in the first tercile. Results for the second and third terciles are summarized in the bottom panel of the same figure. The red line in the top left panel shows the probability of being employed for individuals who transitioned to unemployment in period 0 relative to those who did not. This probability falls to almost zero after the unemployment event, but rises back to more than 60% 12 months after. Note that an individual is counted as employed if they were employed for at least half of the month. In the figure, the only slight drop in employment probability at period 0 is explained by the fact that most individuals transition to unemployment on the last day of the month (i.e., they are observed as employed on the last day of the month, but not on the first of the subsequent month). The steep re-employment slope is likely due to the fact that our sample is comprised of individuals who are closely attached to the labor force. The green line in the same graph shows how an expansionary monetary policy surprise, which causes aggregate earnings to rise by one percent over twelve months, affects the employment probability. The probability rises by about three percentage points one year after the unemployment event. This is in line with the results reported in Figure 3, for more granular sorting. The right panel in Figure 5 isolates the effect of monetary policy. It rises for about 12 months and then stabilizes.

The results for the other terciles are similar, but the effect materializes later after the shock, as evident from the bottom panel in Figure 5. In the second tercile, monetary policy appears to have no significant effects on employment probabilities initially. The effect in the third tercile is more similar to the first, again implying a U-shaped pattern across terciles, similar to the one for earnings seen in Section 4.1. After two years, the effects of monetary policy are not significantly different across terciles.

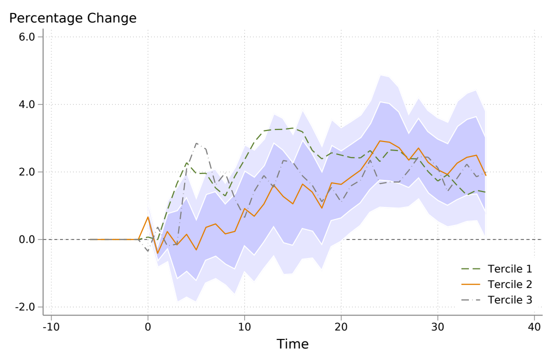
Figure 5: Effect of monetary policy shock on re-employment probabilities




(a) First tercile




(b) First tercile – Difference




(c) Employment

**Note:** The *Top Left Panel* shows the employment probability of individuals in the first tercile who transition into unemployment in month $t = 0$, relative to those who don't (red line). The green line shows how this probability changes after a monetary policy surprise (scaled to be consistent with a one-percent increase in aggregate earnings). The shaded area indicates the 90 percent confidence band based on HAC standard errors. The *Top Right Panel* isolates the effect of a monetary policy surprise on the employment probability, i.e. the difference between the two lines. The *Bottom Panel* shows the same effect for all three terciles. The dashed green lines represent estimates for the first tercile, the solid yellow lines those for the second and the dash-dotted grey line those for the third. The shaded areas in the top right and bottom graph represent 68 and 90 percent confidence bands based on HAC standard errors for the second tercile. Terciles are constructed based on average earnings during the five years prior to the start of the unemployment episode in period 0, conditional on gender and five-year age brackets. The sample period is 2000-2013.

Figure 14 in Appendix B.3 shows the results for the same exercise using the earnings ratio as the dependent variable. The results are very similar, implying that much of the earnings response to monetary policy is in fact driven by employment transitions.

Together, Figures 5 and 14 show that monetary policy indeed has long-run effects on individual earnings that persists substantially beyond the 12-months horizon that the previous sections focus on. Moreover, the effect on employment probabilities also accounts for a large fraction of these long-run earnings-effect of monetary policy. The most pronounced differences in the responses across the income distribution, however, are found at shorter horizons. These conclusions are robust to the inclusion of unemployment benefits into the earnings definition, as we show in the next section.

## 4.5 Accounting for unemployment benefits

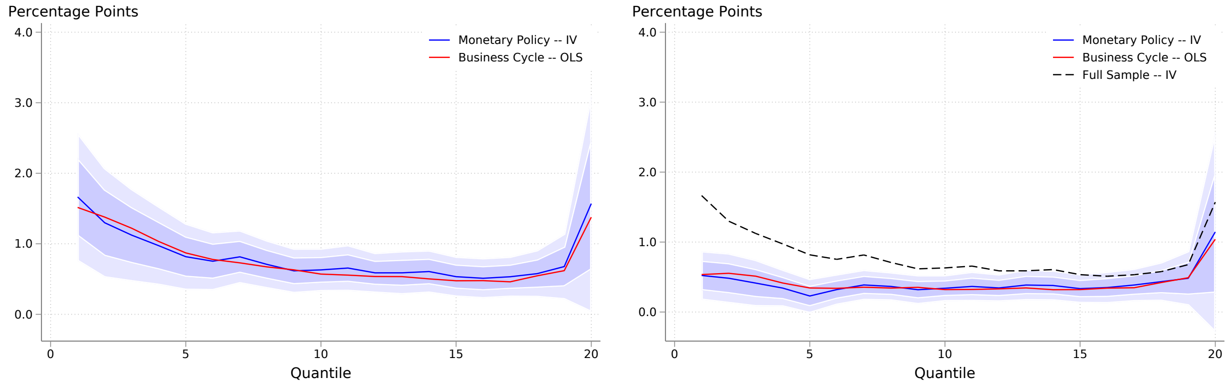
So far, our analysis has concentrated on the effect of monetary policy surprises on labor earnings growth. To that end, we set labor earnings to zero for individuals who are unemployed in period $t$. However, our dataset contains some information about unemployment benefit receipts.[16] In order to come closer to a measure income that is relevant for consumption and savings decisions and thus welfare, in this subsection we thus look at the effect of monetary policy on income including transfers, similar to Amberg et al. (2022). For individuals who are unemployed but do not receive benefits, we set their income to zero. We do not recompute aggregate earnings growth (used to compute the red-lines in the figures), in order to retain comparability with the main results.

Figure 6 reports the results, analogous to Figure 2 in the paper. Monetary policy surprises again affect the bottom of the distribution the strongest, but the effect is somewhat muted. This brings the overall effect (black dotted line in the right panel) closer to the earnings changes of individuals who remain employed. Naturally, the blue and red lines in the right panel remain unchanged, as they do not contain earnings received in unemployment.

Thus including unemployment benefits, to the extent that we have information on them, does not alter our conclusions that (i) the earnings growth of individuals at the bottom of the permanent income distribution is most affected by monetary policy surprises and (ii) the main reason for the heterogeneity is extensive margin transitions.

[16] For more information, see Section A.

Figure 6: Regression coefficients $\beta_{12}^q$ across the income distribution



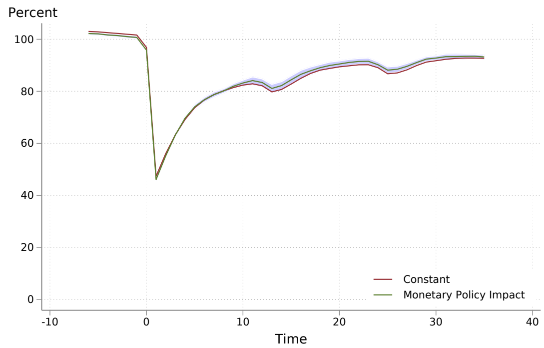

(a) Full sample

(b) Only individuals employed in $t-1$ and $t+12$

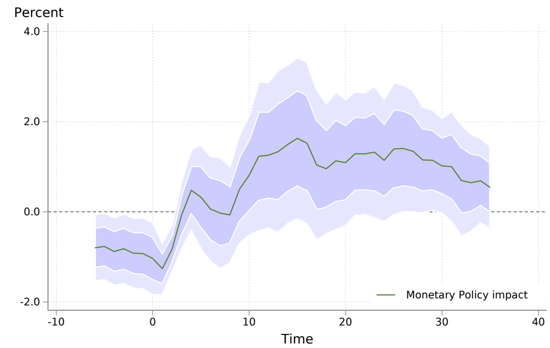
**Note:** The *Left Panel* plots the coefficients $\beta_{12}^q$ in Equation (2) (scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings) and $\beta_{Y,12}^q$ in Equation (3), separately for individuals who shared the same ventile of the permanent income distribution in period $t-1$. Income growth is computed as the log-change in the average income of individuals who were in the same ventile at time $t-1$. In unemployment, earnings are set to unemployment benefits if the individual receives any, otherwise they are set to zero. The *Right Panel* compares the coefficients $\beta_{12}^q$ for the full sample in a ventile (gray dashed line) to $\beta_{12}^{q,E}$ and $\beta_{Y,12}^{q,E}$, estimated on a smaller sample of individuals who are employed both in period $t-1$ and $t+12$ (the blue and red lines, respectively). Ventiles are constructed based on average earnings during the five years prior to $t-1$, conditional on gender and five-year age brackets. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC robust standard errors. The sample period is 2000-2013.

A smilar exercise can be conducted regarding the longer-run earnings effects of unemployment, as in Section 4.2. Specifically, we can substitute the benefit income for the zero-earnings assumption used in that analysis. The top left panel in Figure 7 shows the results of this exercise. The fall in earnings upon unemployment is much less pronounced. Whereas before, earnings fell close to zero upon unemployment, the fall is now closer to 40 percent. However, as the top right panel shows, monetary policy still has a significant impact on earnings. After a year, earnings of individuals who lose their employment during periods of accommodating monetary policy have recovered on average about two percent more than those who become unemployed in normal times. These effects are similar to those for the other two terciles, plotted in the bottom panel of the same figure.

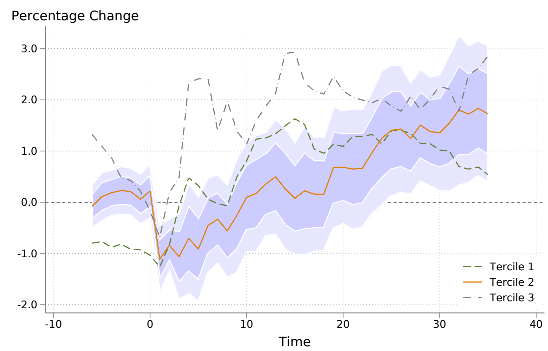
Figure 7: Effect of monetary policy on post-benefit earnings after unemployment




(a) First tercile




(b) First tercile – Difference




(c) Earnings (incl. UI)

**Note:** The *Left Panel* shows the average earnings, including unemployment benefit receipts, of individuals in the first tercile who transition into unemployment in month $t = 0$, relative to those who don't (red line). The green line shows how relative earnings change after a monetary policy surprise (scaled to be consistent with a one-percent increase in aggregate earnings). The shaded area indicates the 90 percent confidence band based on HAC standard errors. The *Right Panel* isolates the effect of a monetary policy surprise on the relative earnings, i.e. the difference between the two lines. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC standard errors. The *Bottom Panel* shows the effect of the same monetary policy surprise on relative earnings including unemployment benefit receipts. The dashed green lines represent estimates for the first tercile, the solid yellow lines those for the second and the dash-dotted grey line those for the third. The shaded areas represent 68 and 90 percent confidence bands based on HAC standard errors for the second tercile. Terciles are constructed based on average earnings during the five years prior to the start of the unemployment episode in period 0, conditional on gender and five-year age brackets. The sample period is 2000-2013.

# 5 Implications for aggregate fluctuations

Recent literature has pointed out that cyclical variations in employment risk may act as an amplifying mechanism for business cycles (Broer et al., 2021; Graves, 2025; Krueger et al., 2016). If workers reduce consumption and build up precautionary savings when separation risk rises in recessions, the resulting contraction in demand could deepen the downturn. In this section, we explore how the heterogeneity we document in Section 4.1 affects the dynamics of aggregate demand in response to monetary policy shocks. Our contribution is to study the heterogeneous incidence of unemployment risk in general equilibrium, and compare the resulting transmission of monetary policy shocks in this setting to previous studies who focused on the heterogeneous incidence of cyclical fluctuations in labor earnings (abstracting from employment risk (Auclert, 2019; Patterson, 2023)) or on cyclical employment risk that is homogeneous across households (Acharya and Dogra, 2020). For this, we study a simple heterogeneous-agent New Keynesian model with both ex-ante and ex-post heterogeneity. Households are ex-ante heterogeneous in their rate of time preference and exposure to unemployment risk. Households are ex-post heterogeneous because of incomplete markets and uninsurable, persistent income and unemployment risk.

## 5.1 Households

The economy is populated by a unit mass of infinitely lived households indexed by $i$. To capture the salient feature of heterogeneous labor-market dynamics across the income distribution, we assume that households belong to one of $g \in \{1, \ldots, G\}$ permanent types with mass $\mu_g$ that differ in labor productivity, labor market risk, and rate of time preference.

### 5.1.1 Preferences

A household $i$ with permanent type $g$ has time-separable preferences over consumption and labor:

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta_g^t \left[ \frac{c_{i,t}^{1-\gamma}}{1-\gamma} - \chi \frac{h_{i,t}^{1+1/\varphi}}{1+1/\varphi} \right], \tag{7}$$

where $c_{i,t}$ is consumption, $h_{i,t}$ are hours worked, $\gamma > 0$ is the coefficient of relative risk aversion, $\varphi > 0$ is the Frisch elasticity of labor supply, and $\chi$ scales the disutility of working.

Households consume, receive income, and can save (but not borrow) by investing in a mutual fund. Hence, they are subject to the following budget constraint:

$$c_{i,t} + a_{i,t+1} = (1 + r_t^p) a_{i,t} + (1 - \tau) y_{i,t} + T_t. \tag{8}$$

where $c_{i,t}$ denotes consumption, $a_{i,t+1} \geq 0$ represents mutual fund holdings, $y_{i,t}$ current labor or replacement income, $\tau$ is a proportional income tax paid by all households, and $T_t$ is a lump-sum transfer that is common across types. The return $r_t^p$ is the ex-post return on the asset holdings of the mutual fund (discussed below), which invests in government bonds and owns the intermediate goods firms.

Households can either be employed $e$ or unemployed $u$. Unemployed households supply zero hours to the market. Transition between employment statuses are described in the next section. Employed households receive labor income, which is proportional to their permanent type-specific productivity level $z_g$, the realization of a persistent productivity shock $x_{i,t}$, the current wage $w_t$, and their hours worked $h_{g,t}$:

$$y^e_{i,t} = w_t h_{g,t} z_g x_{i,t}, \tag{9}$$

When unemployed, households receive unemployment insurance proportional to their steady-state earnings

$$y^u_{i,t} = \theta \overline{w} \overline{h}_g z_g x_{i,t}, \tag{10}$$

where $\theta \in (0, 1)$ is the replacement rate, $\overline{w}$ and $\overline{h}$ are, respectively, the steady levels of wages and hours. We assume that the replacement rate depends on steady state hours and wages to mimic the reality that unemployment benefits depend on a backward-looking measure of labor earnings (proxied here by the steady state).

### 5.1.2 Income Risk

Households face idiosyncratic labor income risk arising from both persistent productivity shocks and employment transitions. Accounting for persistent earnings risk among employed workers, in addition to unemployment risk, is crucial for generating realistic levels of wealth inequality and, consequently, for capturing the consumption response to income shocks, including those stemming from employment transitions (see, e.g., Krueger et al., 2016).

Employed workers face a type-specific job separation rate $\delta_{g,t}$ that varies with the aggregate state. Unemployed households find jobs at type-specific job-finding rates $f_{g,t}$ that also vary with the aggregate state. Employment transitions happen at the beginning of the period before production takes place. The evolution of the type specific unemployment rate is thus given by:

$$u_{g,t} = \delta_{g,t}(1 - u_{g,t-1}) + (1 - f_{g,t})u_{g,t-1} \tag{11}$$

Each household's labor productivity is subject to a persistent idiosyncratic shock $x_t \in \mathcal{X}$ which evolves according to an AR(1) process:

$$\log x_{t+1} = \rho \log x_t + \xi_{t+1}, \quad \xi_{t+1} \sim \mathcal{N}(0, \sigma^2), \tag{12}$$

For parsimony, the persistence and volatility parameters $\rho$ and $\sigma$ are common across all permanent types $g$.

## 5.2 Firms

The formulation of the supply-side of the economy closely follows the New Keynesian literature. A representative competitive firm produces a final good using a continuum of intermediate goods:

$$Y_t = \left( \int_0^1 Y_{j,t}^{\frac{\varepsilon-1}{\varepsilon}} dj \right)^{\frac{\varepsilon}{\varepsilon-1}}, \tag{13}$$

where $\varepsilon > 1$ is the elasticity of substitution.

Taking the level of aggregate demand $Y_t$, cost minimization for the final goods producer implies that the demand for intermediate good $j$ is given by

$$y_{jt} = y(p_{j,t}; P_t, Y_t) = \left( \frac{p_{j,t}}{P_t} \right)^{-\epsilon} Y_t, \tag{14}$$

where $P_t$ is the price of the final good, which, given our CES assumption can be calculated as:

$$P_t = \left( \int_0^1 p_{j,t}^{1-\epsilon} dj \right)^{\frac{1}{1-\epsilon}}.$$

### 5.2.1 Intermediate Goods Firms

Each intermediate firm $j$ produces output linearly using labor as the only input:

$$Y_{j,t} = H_{j,t}. \tag{15}$$

Each firm rents labor in a perfectly competitive market giving rise to real marginal cost $mc_{j,t} = w_t$. Intermediate-goods firms face quadratic adjustment costs in the style of Rotemberg, giving rise to a standard New Keynesian Philips Curve. Each intermediate goods firm faces a constant operating cost $F$ which we use to set profits to zero in steady state. Firms are owned by the mutual fund which collects the profits of the intermediate goods firms.

## 5.3 Mutual fund

All household savings are collected by a mutual fund, which invests them into government bonds. The mutual fund also owns the intermediate goods firms and distributes their profits in proportion to household savings. That is, $r_t^p = r_t + \pi_t / A_{t-1}$, where $\pi_t$ are the profits from intermediate goods firms, $r_t$ is the return on government debt, and $A_{t-1}$ are total households assets collected in period $t-1$.

## 5.4 Government

The government consumes a fixed amount $G_t = \bar{G}$, issues real debt $B_t$, levies taxes on labor income $W_t H_t$ and unemployment benefits, and pays unemployment insurance benefits $UI_t$ and lump-sum transfers $T_t$. The period budget constraint is:

$$B_t = (1 + r_t)B_{t-1} + G_t + (1 - \tau)UI_t + T_t - \tau W_t H_t \tag{16}$$

Transfers adjust to ensure budget balance in every period. The central bank follows a Taylor-type interest rate rule:

$$i_t = \bar{i} + \phi_\pi \pi_t + \varepsilon_t^{mp}, \tag{17}$$

where $\varepsilon_t^{mp}$ is an exogenous monetary policy shock.

## 5.5 Equilibrium

An equilibrium consists of sequences of prices, allocations, policy functions, and distribution such that households and firms optimize, government policies satisfy the budget constraint, monetary policy follows its rule, markets clear, and the law of motion for the household distribution is consistent with individual decisions.

Goods market clearing requires:

$$Y_t = C_t + G_t - F. \tag{18}$$

Asset market clearing requires:

$$\int a_{i,t+1} \Omega_t(di) = A_t = B_t. \tag{19}$$

Consistency of the household distribution requires:

$$\Omega_{t+1} = \mathcal{H}_t(\Omega_t) \tag{20}$$

where $\mathcal{H}_t$ is generated by the policy functions $a_{t+1}$ and stochastic processes for $x$ and $e$.

## 5.6 Calibration

We first calibrate a set of parameters to target empirical moments from aggregate data and the literature. Then, our calibration chooses a second set of parameters to target key features of the German micro-data.

The model is quarterly. We choose a coefficient of relative risk aversion of 2, and an elasticity of substitution across intermediate goods equal to 11, compatible with an average markup of 10% (Born and Pfeifer, 2014). We calculate the real interest rate in the data by subtracting the inflation rate from the the Eonia interbank rate, yielding an average yearly real interest rate of 1.231%. Consequently, we set $1 + r = 1.01231^{\frac{1}{4}}$. For our sample period, German final government consumption as a share of GDP was 19.16%, which we use to discipline government spending $G$ in the model (Eurostat, 2026). We set the slope of the price Phillips curve to $\kappa_p = 0.10$, following Beschin et al. (2025).

Regarding the persistent earnings risk of the employed, which is homogeneous across all households, we assume that the process has a quarterly persistence of $\rho = 0.9695^{\frac{1}{4}} = 0.9923$ and standard deviation of $\sigma = 0.098$. These values are from Krueger et al. (2016) and estimated on disposable income conditional on employment using the PSID in the US. Krueger et al. (2010) show that the (persistent component) of the earnings process in Germany is similar to that in the US, justifying our use of the US numbers. We transform this continuous stochastic process to a finite-state Markov chain with three income states using the Rouwenhorst (1995) method. These parameters are summarized in Table 2.

Table 2: Externally calibrated parameters

| **Description** | **Parameter** | **Value** | **Source** |
|---|---|---|---|
| Risk aversion | $\gamma$ | 2.00 | Standard |
| Phillips Curve slope | $\kappa_p$ | 0.10 | Beschin et al. (2025) |
| Replacement rate | $\theta$ | 0.60 | Data |
| Income shock persistence | $\rho$ | 0.99 | Krueger et al. (2016) |
| Income shock variance | $\sigma$ | 0.10 | Krueger et al. (2016) |
| Government spending | $G/Y$ | 0.19 | Data |
| Real interest rate in pct. | $r$ | 0.31 | Data |
| Elasticity of substitution | $\varepsilon$ | 11.00 | Born and Pfeifer (2014) |

**Note:** The Table summarizes the values chosen for model parameters which are externally calibrated and unrelated to labor market transition risk.

To capture the key features of heterogeneous earnings risk across the income distribution while keeping the model computationally manageable, we set the number of permanent types

$G = 3$ and let $g \in \{1, 2, 3\}$. Types 1 and 3 represent, respectively, the bottom and top 30 percent of the income distribution. Type 2 represents the middle 40 percent. We normalize the productivity of each group to match the group's average earnings (while employed). We calibrate the steady-state separation ($\delta_{ss}^g$) and job-finding ($f_{ss}^g$) using the average type-specific values in the German labor histories data.

Table 5 summarizes the results of this exercise. The reported mean earnings are displayed in the first row of Table 5. As in Table 1, earnings are strongly increasing along the permanent income distribution. Similarly, the three-month separation probability is decreasing from around 3 percent for Type 1 to less than 1 percent for Type 3, while the three-month job-finding probability follows an inverted U-shape. Low types thus face significantly higher rates of employment risk and receive lower earnings while employed. In steady state, the type-1 agents have an unemployment rate of almost 18% compared to just over 3% for the type-3 agents.

As is well-known from the HANK literature, the average degree of consumption insurance (as summarised by the average marginal propensity to consume out of transitory income shocks) is a key determinant of policy transmission. Consistent with quasi-experimental evidence and structural interpretations of household liquidity constraints, we assume quarterly MPCs of 0.30, 0.25, and 0.20 across three permanent types, reflecting the well-documented decline in consumption responses with income and liquid wealth (Fagereng et al., 2021; Johnson et al., 2006; Kaplan et al., 2014; Parker et al., 2013). We target these values, together with the real interest rate, by choosing the type-specific discount factors $\beta^g$ and the steady-state level of government bonds $B$.

Table 3: Internally calibrated parameters

| | $B/Y$ | $\beta_1$ | $\beta_2$ | $\beta_3$ |
|---|---|---|---|---|
| Heterogeneous risk | 3.93 | 0.9523 | 0.9766 | 0.9833 |
| Homogeneous risk | 2.59 | 0.9683 | 0.9756 | 0.9772 |

**Note:** The Table depicts the Asset-to-output ratio $B/Y$ (Column 1), and the type-specific discount factors ($\beta_i$ $i = 1, 2, 3$, Columns 2 to 4) in the benchmark model with heterogeneous unemployment risk, and an alternative model with homogeneous unemployment risk across types.

The first row of Table 3 summarizes the internally calibrated parameters. To match the calibration targets, the benchmark economy with heterogeneous risk requires an asset-to-income ratio of close to 4. Furthermore, the discount factors are increasing throughout the type distribution.

Table 4 compares the model-implied wealth distribution to one constructed using the ECB's distributional wealth accounts for Germany (European Central Bank, 2026). The empirical average Gini coefficient for wealth from 2011q1 to 2025q2 is 0.77, and the model matches it very closely. Similarly, the model accurately matches the wealth held by different shares of the distribution.

Table 4: Wealth distribution

| | Data | Het. risk | Hom. risk |
|---|---|---|---|
| Wealth Gini | 0.77 | 0.80 | 0.74 |
| Bottom 50% share (%) | 2.2 | 1.4 | 2.6 |
| Decile 6 share (%) | 3.8 | 1.5 | 2.4 |
| Decile 7 share (%) | 6.5 | 3.4 | 4.2 |
| Decile 8 share (%) | 10.5 | 7.1 | 11.1 |
| Decile 9 share (%) | 16.7 | 22.6 | 26.1 |
| Decile 10 share (%) | 60.3 | 64.0 | 53.6 |

**Note:** Data on wealth shares come from ECB's distributional wealth accounts. It is constructed as the share of total net household wealth held by each part of the net wealth distribution. Model wealth shares are computed for both the heterogeneous and homogeneous risk steady states.

## 5.7 Aggregate risk

The aggregate shock that we consider in the model is an innovation to the Taylor rule, eq. (17). We assume that $\epsilon_t^{mp}$ follows an AR(1) with persistence of 0.8. Following Boppart et al. (2018), we linearize the model in sequence space using MIT shocks. Since we are solving the model to first order, we posit a linear relationship between the cyclical risk from movements in separation and job-finding rates and aggregate activity. Specifically, the ype-specific employment transition probabilities are given by

$$\delta^g(w) = \delta^g_{ss} + \delta^g_1 \left(\log(w) - \log(w_{ss})\right) \tag{21}$$

$$f^g(w) = f^g_{ss} + f^g_1 \left(\log(w) - \log(w_{ss})\right) \tag{22}$$

where $\delta^g$ is the probability of a worker of permanent type $g$ transitioning from employment to unemployment, and $f^g$ is the type-specific job-finding probability.[17] The cyclicality of

[17] This specification is analogous to a wage Phillips curve, linking labor market slack to wage dynamics in a standard way. It also resolves a potential overdeterminacy problem in the model if transitions depend on aggregate earnings.

each depends on the contemporaneous realization of the aggregate wage according to the parameters $\delta_1^g$ and $f_1^g$. This formulation allows us to match, in a parsimonious way, the heterogeneous employment transition probabilities we document in the data, both in the steady state and along the transition. Relatedly, in our benchmark economy, we hold hours of the employed fixed, so the sole margin of labor adjustment is through hiring and separations. This parsimonious choice reflects both modeling pragmatism and empirical relevance: our estimates in Section 4.1 show that earnings growth conditional on continuous employment responds only modestly to monetary policy, with little heterogeneity across the income distribution.

Table 5: External parameters governing income risk

| | (1)<br>Type 1<br>mean | (2)<br>Type 2<br>mean | (3)<br>Type 3<br>mean |
|---|---|---|---|
| Earnings (employed) | 1709.84 | 2558.71 | 3869.66 |
| Separation probability ($\delta_{ss}^g$) | 3.25 | 1.23 | 0.71 |
| Cyclicality of separations ($\delta_1^g$) | -1.45 | -0.53 | -0.28 |
| St. err. separations | 0.46 | 0.17 | 0.09 |
| Job-finding probability ($f_{ss}^g$) | 14.89 | 23.00 | 20.53 |
| Cyclicality of job-finding ($f_1^g$) | 5.38 | 5.80 | 6.35 |
| St. err. job-finding | 1.60 | 1.77 | 2.11 |

**Note:** The Table shows the model-relevant empirical results for the three income groups. Earnings of the employed are expressed in 2010 euros. The separation and job-finding probabilities represent the probability of transitioning between months $t-1$ and $t+3$ to approximate quarterly transitions. They are expressed in percent. The cyclicality estimates are computed according to Equation (5) and are expressed in percentage points; we also report the associated standard errors for the cyclicality estimates. The sample period is 2000-2013.

To discipline the parameters in Equations (21) and (22), i.e., the type-specific transition probabilities and their cyclicalities, we use the estimates from Equation (5), estimated for each of our three types (i.e., the bottom and top 30%, and the middle 40% of the income distribution). The estimates are reported in figure 3. Importantly, Table 6 shows that when we estimate the corresponding elasticities of average individual earnings for the three parts of the income distribution, the model matches these well.

## 5.8 Results

Figure 8 shows the baseline model's response to a 25 basis point increase in the central bank's policy rate. Consumption falls upon impact by approximately 1.3 percent, driving

Table 6: Type-specific earnings elasticities

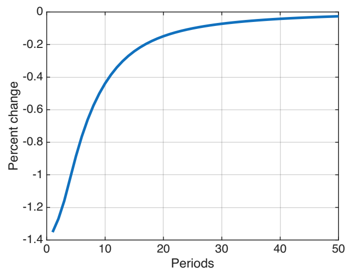
| Type | Data elasticity | Model elasticity |
|---|---|---|
| 1 | 1.455 | 1.587 |
| | (0.552) | |
| 2 | 0.838 | 0.976 |
| | (0.250) | |
| 3 | 0.901 | 0.785 |
| | (0.366) | |

**Note:** The Table depicts the earnings elasticities, estimated according to Equation (2) for three types, in the model and the data. As the model is quarterly, we use the four-quarter ahead elasticity to compute the model elasticities, while we use h = 12 in the monthly data. For the data, standard errors are reported in parentheses.

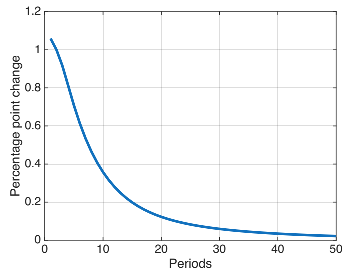
an increase in unemployment. For this to be consistent with the earnings-unemployment-transition relationship (21) and (22), the real wage must fall. Accordingly, the probability of an employed worker moving into unemployment increases, while the job-finding probability of the unemployed decreases. Unemployment rises by slightly more than one percentage point at its peak before gradually returning to steady state.

The baseline results show that monetary policy has substantial effects on consumption and unemployment in our benchmark economy with heterogeneous incidence of unemployment risk. To understand the mechanisms driving these responses, and to quantify the importance of both unemployment risk and its heterogeneous incidence, we now compare the benchmark to counterfactual economies. We proceed in two steps. First, we ask whether it matters that labor adjustment occurs through employment transitions rather than hours worked by the employed. Second, we ask whether the concentration of unemployment risk among low-income workers amplifies monetary transmission.

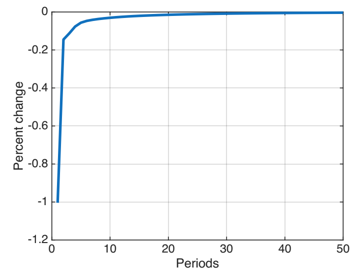
Figure 8: Impulse responses with heterogeneous unemployment incidence




(a) Consumption




(b) Unemployment



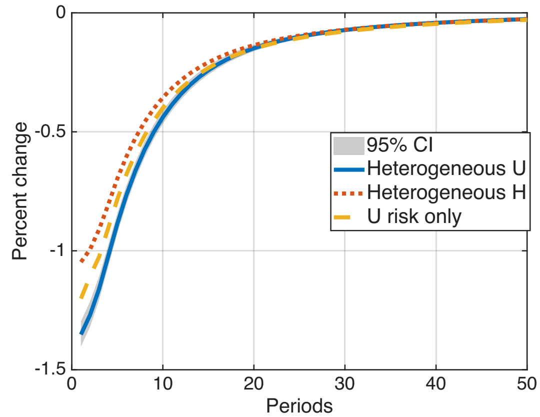

(c) Real wage

**Note:** The Figure shows the model-generated impulse responses to a 25 basis point monetary policy contraction for our benchmark economy. Unemployment incidence is heterogeneous by permanent type in the steady state. Furthermore, along the transition, labor market transition probabilities move heterogeneously across types, governed by Equations (21) and (22). The *Left Panel* shows the consumption impulse response, the *Middle Panel* shows impulse responses of unemployment, the *Right Panel* shows the impulse responses of the real wage.

## 5.9 The Role of Unemployment Risk

Our first counterfactual, which we label "Heterogeneous H", departs from the same steady state as the benchmark economy, with heterogeneous unemployment risk across types. However, in response to a monetary policy shock, type-specific hours worked by the employed adjust while unemployment remains constant. For comparability, we impose the same type-specific real-wage-labor-input schedule as in the benchmark. This means that, for a given change in the real wage, type-specific hours movements imply the same change in type-specific labor inputs (and hence aggregate productivity) as movements in type-specific unemployment in the benchmark model. Thus, shocks still affect types differently, but fluctuations in type-specific earnings arise from hours rather than employment transitions. This counterfactual is similar in spirit to Patterson (2023), where incidence is heterogeneous across agents but not modeled as unemployment risk.

Figure 9: Consumption response to monetary policy shock




**Note:** The Figure shows the model-generated consumption impulse responses to a 25 basis point monetary policy contraction for three different cases: the benchmark economy (Heterogeneous U, solid blue), an economy with unequal incidence of unemployment across types in steady state and heterogeneous hours adjustments across types along the transition (Heterogeneous H, dashed red), and a partial-equilibrium counterfactual using benchmark policy functions but Heterogeneous H distributions along the transition, to isolate the role of anticipated but not realized unemployment risk (U risk only, dotted yellow). The shaded area represents the 95% confidence interval for the Heterogeneous U response.

Figure 9 plots the consumption response in the benchmark economy (Heterogeneous U, solid blue) together with that in the Heterogeneous H counterfactual (dashed red). When heterogeneous earnings changes materialize through hours worked rather than unemployment, consumption falls by close to one percent on impact—approximately 23 percent less than in the benchmark economy. The cumulative difference between the consumption responses, reported in Table 9, is similar, equal to 16 percent.

The additional amplification in the benchmark model is driven by two forces. First, unemployment risk is higher in the Heterogeneous U economy due to a rise in separation probabilities and a fall in job-finding rates. This increases precautionary savings, leading to a larger fall in aggregate consumption demand. We call this the *precautionary-savings channel*. Second, because in the benchmark economy the realized decline in average earnings is concentrated among households that become unemployed, it implies a larger decline in consumption demand: the average consumption effect of unemployment shocks is larger than that of a more widespread but smaller earnings drop from a homogeneous fall in hours. We call this the *realized-income channel*.

To quantify the relative importance of these two channels, we construct an intermediate impulse response analogous to an Oaxaca-Blinder decomposition. In this partial-equilibrium exercise, we use the consumption-savings policy functions from the Heterogeneous U economy, which include the precautionary-savings response to increased unemployment risk, but the distribution of households from the Heterogeneous H economy, which lacks the realized unemployment shocks:

$$C_t^{risk} = \int c_t^{HetU}(s)\Omega_t^{HetH}(ds). \tag{23}$$

where $s$ denotes the idiosyncratic state $s$ (consisting of permanent type, asset holdings, earnings level and employment state), and $\Omega_t^{HetH}(s)$ denotes the distribution across states in the Heterogeneous H economy. The yellow dotted line in Figure 9 plots the corresponding impulse response. It lies between the two original counterfactuals, suggesting that unemployment risk induces a substantial precautionary-savings response. The decomposition indicates that roughly 50 percent of the additional consumption response in the benchmark economy is driven by increased precautionary savings; the remaining 50 percent reflects the larger consumption impact of realized unemployment shocks.[18]

Table 7: Propensities to consume

| | Heterogeneous Incidence | | Homogeneous Incidence | |
|---|---|---|---|---|
| Type | MPC | ACES | MPC | ACES |
| Type 1 | 0.30 | 0.71 | 0.30 | 0.60 |
| Type 2 | 0.25 | 0.45 | 0.25 | 0.47 |
| Type 3 | 0.20 | 0.33 | 0.20 | 0.45 |

**Note:** The table displays the marginal propensities to consume and the consumption response to unemployment, ACES, for the homogeneous and heterogeneous steady states of the economy. Consumption response to unemployment is expressed as the change in consumption in the first period of unemployment, relative to the income fall in the first period.

Table 7 illustrates the realized-income channel more directly. It reports, for the benchmark steady state, both standard marginal propensities to consume out of small one-time transfers (MPC) and the average consumption effect of a separation shock (ACES), defined as the consumption change in the first period of unemployment relative to the income loss.[19] A type-1 (low-income) worker who becomes unemployed—losing roughly 40 percent of her after-tax income—reduces consumption by about 70 cents for every dollar of lost income. A

[18]The shaded area in the figure represents the 95% confidence interval for the Heterogeneous U economy, which we construct by bootstraping $\delta_1^g$ and $f_1^g$ 50 times, using the point estimates and standard errors reported in Table 5.

[19]ACES is the analogue to an MPC for a separation into unemployment. It takes into account that separation involves a discrete change in income (as opposed to an MPC being a marginal change) as discussed in Auclert and Mitman (2022, 2023) who study the consumption effects of consumer default.

type-3 (high-income) worker, in contrast, reduces consumption by less than half as much, as her unemployment is less persistent and her savings buffer is larger. Importantly, propensities to consume out of unemployment shocks are substantially more dispersed than standard MPCs, and substantially more dispersed than the consumption effects of persistent hours changes in the Heterogeneous H counterfactual. This explains the additional amplification from unemployment risk in the benchmark economy.

### 5.10 The Role of Heterogeneous Incidence

The previous comparison held fixed the heterogeneous incidence of unemployment risk across the income distribution. We now ask: does the concentration of unemployment risk among low-income, high-MPC workers amplify monetary transmission beyond what would obtain if risk were distributed equally?

Table 8: External parameters governing income risk

| | Single Type mean |
|---|---|
| Earnings (employed) | 2721.88 |
| Separation probability ($d_0^g$) | 1.52 |
| Cyclicality of separations ($d_1^g$) | -0.63 [-0.56] |
| St. err. separations | 0.20 |
| Job-finding probability ($f_0^g$) | 17.25 |
| Cyclicality of job-finding ($f_1^g$) | 5.61 [4.54] |
| St. err. job-finding | 1.69 |

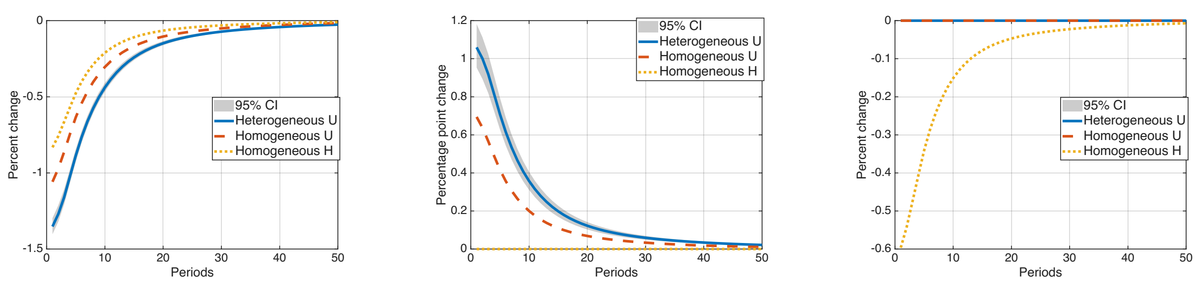
**Note:** The Table shows the same model-relevant empirical results as in Table 5 but for the full sample. Earnings of the employed are expressed in 2010 euros. The separation and job-finding probabilities represent the probability of transitioning between months $t-1$ and $t+3$ to approximate quarterly transitions. They are expressed in percent. The cyclicality estimates are computed according to Equation (5) and are expressed in percentage points; we also report the associated standard errors for the cyclicality estimates. The terms in square brackets are the cyclicalities used in the model counterfactuals with homogeneous cyclicality. The sample period is 2000-2013.

To answer this question, we construct two additional counterfactual economies with homogeneous unemployment incidence. In these economies, all household types face identical employment transition probabilities in the steady state. Table 8 reports the relevant empirical moments, estimated on the full sample rather than by type. We recalibrate the type-specific discount factors to achieve the same type-specific MPCs as the benchmark economy; the resulting parameters are reported in the second row of Table 3. The economy with homogeneous risk requires an asset-to-output ratio of 2.6, compared to 4.0 in the benchmark. This decrease arises because homogeneous risk loads unemployment more heavily on high-income workers: separation probabilities are higher and job-finding probabilities are lower in Table 8 compared to those for type-3 workers in Table 5. Thus, high-income agents have both lower income (due to more unemployment) and stronger incentives to save. To achieve the same MPC distribution, the discount factor for type-3 households must fall, reducing aggregate saving.

In our "Homogeneous U" counterfactual, households face homogeneous unemployment risk both in steady state and in response to shocks. We calibrate the cyclicality of transition rates to ensure that aggregate labor input responds identically to real wage changes as in the benchmark; this requires using productivity-weighted averages of the type-specific cyclicality estimates rather than the empirical full-sample estimates. Table 8 reports these coefficients in square brackets. This adjustment ensures that the supply side of the economy—specifically, the relationship between real wages and aggregate labor input—is unchanged across specifications. The economy with homogeneous unemployment risk and employment adjustments along the transition is similar to the ones in Ravn and Sterk (2021) and Graves (2025): a decrease in output increases unemployment risk across the economy.[20]

Our final counterfactual, "Homogeneous H," combines homogeneous unemployment risk in steady state with hours adjustment along the transition, providing a natural comparison to an economy without either heterogeneous incidence or extensive-margin adjustment.

Figure 10: Heterogeneous and homogeneous incidence




(a) Consumption (b) Unemployment (c) Hours

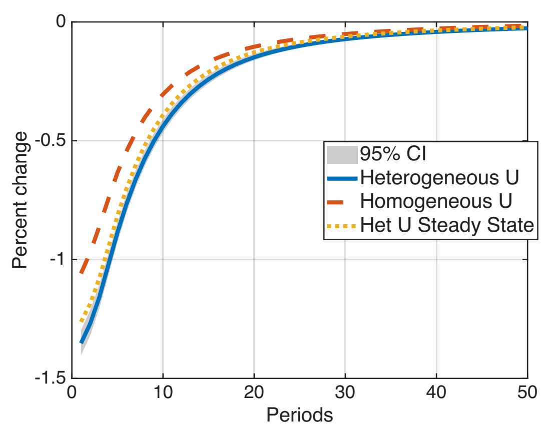
**Note:** The Figure shows the model generated impulse responses to a 25 basis point monetary policy contraction for three economies: the benchmark economy (Heterogeneous U, solid blue), an economy with equal incidence of unemployment in steady state and along the transition (Homogeneous U, dashed red), and an economy with equal unemployment incidence in steady state, but hours adjustments along the transition (Homogeneous H, dotted yellow). The shaded area represents the 95% confidence interval for the Heterogeneous U response. The *Left Panel* shows the consumption impulse responses, the *Middle Panel* shows impulse responses of unemployment, the *Right Panel* shows the impulse responses of hours worked by the employed.

[20]In Graves (2025), the rise in unemployment is due to a decrease in labor demand paired with sticky wages, in our model, the relationship between output and unemployment is exogenously given by our empirical estimates.

Figure 10 plots the impulse responses for all three economies. The left panel shows that the Homogeneous U economy (red dashed line) responds substantially less to the same monetary policy shock than the benchmark. The impact response is approximately 25 percent weaker. This suggests that the distribution of unemployment incidence across permanent types has a meaningful impact on the response of aggregate variables to shocks.

Part of this weaker response is visible in the middle panel: unemployment rises by only 0.7 percentage points in the Homogeneous U economy, compared to 1.1 percentage points in the benchmark. Since homogeneous incidence loads unemployment more heavily on highly productive workers, a smaller increase in aggregate unemployment is required to achieve the same fall in aggregate output. However, this compositional effect is not the only force at play. Even conditional on the same aggregate output response, heterogeneous incidence amplifies because it concentrates unemployment risk on workers who have higher MPCs and account for the bulk of precautionary savings.

Figure 11: Same steady state




**Note:** The Figure shows the model-generated consumption impulse responses to a 25 basis point monetary policy contraction for three different cases: the benchmark economy (Heterogeneous U, solid blue), an economy with equal incidence of unemployment across types (Homogeneous U, dashed red), and an economy with unequal incidence of unemployment in steady state, but homogeneous unemployment-rate movements along the transition (Het U Steady State, dotted yellow). The shaded area represents the 95% confidence interval for the Heterogeneous U response.

To isolate the role of heterogeneous incidence along the transition while holding the steady state fixed, Figure 11 compares the benchmark to an economy with heterogeneous steady-state unemployment risk but homogeneous unemployment movements in response to shocks (yellow dotted line). This intermediate case shows that heterogeneous incidence during the transition accounts for approximately one third of the difference between the Heterogeneous U and Homogeneous U economies.

Table 9: Impulse response comparison

| | Impact Response | Cumulative Response |
|---|---|---|
| Heterogeneous $U$ (Benchmark) | 1.000 | 1.000 |
| Heterogeneous $H$ | 0.775 | 0.834 |
| Homogeneous $U$ | 0.783 | 0.703 |
| Homogeneous $H$ | 0.616 | 0.502 |

**Note:** The table computes the impulse response sizes of four counterfactual economies, relative to the benchmark economy (Heterogeneous U). Heterogeneous H is an economy with heterogeneous type-specific unemployment probabilities in steady state, in which hours adjust along the transition. Homogeneous U and Homogeneous H have homogeneous type-specific unemployment probabilities in steady state, and all types are affected homogeneously by unemployment and hours adjustments, respectively, along the transition. The impact response is calculated as the initial log-deviation from steady state. The cumulative response is calculated as the cumulative sum of all log-deviations from steady state along the entire transition.

Table 9 summarizes the results across all four specifications (the full impulse responses to the four economies are plotted in Figure 20 in the Appendix). Relative to the benchmark, the Homogeneous H economy—with neither unemployment risk nor heterogeneous incidence—exhibits an impact consumption response that is 38 percent smaller and a cumulative response that is 50 percent smaller. The comparison between Heterogeneous U and Homogeneous U isolates the role of heterogeneous incidence, while the comparison between Homogeneous U and Homogeneous H isolates the role of unemployment risk per se under equal incidence. Both margins contribute meaningfully to amplification, with their combined effect being roughly multiplicative.

# 6 Conclusion

This paper makes two contributions to our understanding of monetary policy transmission. First, using high-frequency German administrative data, we document that monetary policy surprises affect earnings growth substantially more at the bottom of the income distribution than at the top. Earnings in the bottom quintile respond roughly three times as strongly as earnings at the median. This heterogeneous incidence is driven almost entirely by the extensive margin: separation rates into unemployment are far more countercyclical for low-income workers, while job-finding rates and earnings growth conditional on employment respond relatively uniformly across the distribution. These patterns are not specific to monetary policy—they appear similarly in the unconditional comovement of individual and aggregate earnings—suggesting they reflect a general feature of labor market dynamics.

Second, we develop a heterogeneous-agent New Keynesian model to quantify how this unequal incidence shapes the aggregate response to monetary policy. We show that cyclical unemployment risk amplifies consumption responses through two channels: a precautionary-savings channel, as workers reduce consumption when separation risk rises, and a realized-income channel, as a rise in unemployment reduces consumption by more than an equivalent reduction in earnings from hours worked. Our decomposition suggests these channels are roughly equally important. Heterogeneous incidence further amplifies transmission by concentrating unemployment risk on low-income, high-MPC workers who account for the bulk of precautionary savings. Taken together, these forces make the consumption response to monetary policy roughly 100 percent larger than in an economy with homogeneous hours adjustment.

While our empirical findings are for Germany, we believe they are likely applicable more broadly. The heterogeneous incidence of earnings risk has been documented across numerous advanced economies, and the mechanisms we identify—countercyclical separation risk concentrated among low-income workers—are consistent with search-theoretic models of the labor market that are not specific to any institutional setting.

Our findings carry implications for both research and policy. For the HANK literature, our results suggest that explicitly modeling unemployment risk and its heterogeneous incidence is quantitatively important for understanding monetary transmission. Models that capture heterogeneous earnings responses solely through hours or productivity differences may understate both the level and the distribution of consumption responses to policy. For policymakers, our results highlight that monetary expansions reduce income inequality by disproportionately improving labor market outcomes for low-income workers, but also that recessions impose particularly severe costs on this group. Policies that reduce the cyclicality of separation risk for vulnerable workers, through employment protection, automatic stabilizers, or targeted interventions, could potentially dampen aggregate fluctuations by muting the precautionary-savings channel.

Understanding the mechanisms by which monetary policy generates heterogeneous labor-market outcomes is an important next step. Moser et al. (2021) make progress in this direction using similar administrative data, studying how interest rate changes affect credit supply to firms and, through this channel, the employment prospects of their workers. Designing and evaluating policies that target the underlying sources of heterogeneous unemployment risk is a promising avenue for future research.

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Seasonal adjustment: Seasonally and calendar adjusted data.
Unit of measure: Index (2010 = 100).
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Time frequency: Monthly.
Seasonal adjustment: Seasonally adjusted data, not calendar adjusted.
Age class: Total.
Sex: Total.
Unit of measure: Percentage of population in the labour force.
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Sector: General government.
Unit of measure: Percent of GDP.
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# Appendix

## A Micro data

We use the Sample of Integrated Labor Market Biographies (SIAB) data. The SIAB data is provided in the form of labor market spells, each at most one year in duration, reporting the average daily wage during the spell. We convert these spells into monthly observations and multiply the daily wages by 30 in order to ascertain monthly earnings. If an individual reports multiple simultaneous spells during a month, we keep the spell that is classified as "Subject to social security without special characteristics" (as classified in Table A4 of Ganzer et al. (2017)). If one of the simultaneous spells implies non-employment, we keep that spell and classify the individual as non-employed. We classify individuals who earn less than the lower social security contribution limit as non-employed. All non-employed workers are coded to have zero income.

We classify as unemployed those individuals who receive unemployment benefits (ALG I). Because the definition and eligibility of these benefits changed over time, we declare any individuals who are non-employed but started their non-employment spell in unemployment as unemployed for the whole duration of the non-employment spell. For these individuals, we substitute previous levels of unemployment benefits for the periods in which these values are missing. This addresses in particular the shortening of unemployment benefit eligibility around 2005. We do not use data on receipts of "Arbeitslosenhilfe", i.e., benefits paid to individuals who were unemployed for too long, thus becoming ineligible for unemployment benefits. Regulations regarding these payments changed significantly in 2005; for consistency, we exclude them. All earnings are deflated into real earnings using the monthly CPI index (OECD, 2025).

# B Additional results

## B.1 Alternative measure of permanent income

In our baseline estimation, we sort individuals into quantiles based on their average earnings over the five year period preceding the monetary policy shock. This is our preferred measure of permanent income. A potential concern with this approach is that a five-year period is not long enough to accurately proxy for permanent income. In this case, some of the variation in our permanent income measure would be driven by temporary income fluctuations. In this section, we reproduce our main results for an alternative measure of permanent income: the fixed effect in a Mincer regression.

We estimate the following specification for all employed individuals in our sample:

$$\log(earn_{i,t}) = \alpha + \lambda_i + \beta_{ea} age_t + \beta_{eb} age_t^2 + \varepsilon_{i,t,ea} \tag{24}$$

where $age_t$ and $age^2$ represent a second order polynomial of individual $i$'s age and $\lambda_i$ is the earnings fixed effect for individual $i$, our proxy for lifetime earnings. Subsequently, each month, we sort all individuals into twenty quantiles based this measure.[21]

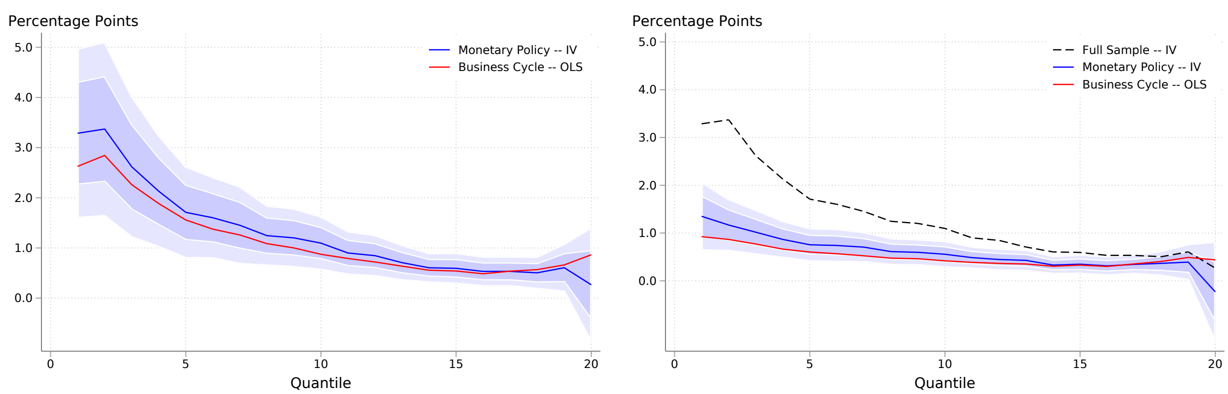
Figure 12 reports the results from this exercise, similarly to 2 in Section 4.1 of the paper. As in the baseline specification, the monetary policy surprise consistent with raising

[21] When estimating the regression in Equation 24, we include all employed individuals in our dataset, including those whose earnings exceed the social security contribution cutoff. However, when computing the effect of monetary policy surprises on earnings growth, we again exclude the censored earnings observations.

aggregate earnings by 1 percent affects the bottom end of the permanent income distribution significantly more than the top. If anything, the downward slope is slightly more pronounced in Figure 12. The increase in the effect at the top of the distribution observed in Figure 2, in contrast, is absent here. As before, the results appear to be driven by extensive margin transitions: the effect of monetary policy shocks on earnings is essentially homogeneous across the income distribution for individuals who remain employed. The effect of monetary policy surprises on labor market transitions (not reported here) are similarly close to those reported in Figure 3.

From the analysis in this section we conclude that our results are robust to the use of other measures of permanent income.

Figure 12: Regression coefficients $\beta_{12}^{q}$ across the income distribution




(a) Full sample

(b) Only individuals employed in $t-1$ and $t+12$

**Note:** The *Left Panel* plots the coefficients $\beta_{12}^{q}$ in Equation (2) (scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings) and $\beta_{Y,12}^{q}$ in Equation (3), separately for individuals who shared the same ventile of the permanent income distribution in period $t-1$. Income growth is computed as the log-change in the average income of individuals who were in the same ventile at time $t-1$. The *Right Panel* compares the coefficients $\beta_{12}^{q}$ for the full sample in a ventile (gray dashed line) to $\beta_{12}^{q,E}$ and $\beta_{Y,12}^{q,E}$, estimated on a smaller sample of individuals who are employed both in period $t-1$ and $t+12$ (the blue and red lines, respectively). **Ventiles are constructed based on the individual fixed effect $\lambda_i$ in regression 24.**The shaded areas indicate 68 and 90 percent confidence intervals based on HAC robust standard errors. The sample period is 2000-2013.

## B.2 Results using different measures of monetary policy surprises

A potential shortcoming of the monetary policy surprises reported in Almgren et al. (2022) is that they may conflate two different components of monetary policy announcements. Nakamura and Steinsson (2018) and Jarociński and Karadi (2020) argue that a larger than expected interest rate cut, for example, may not only be an expansionary monetary policy action, but a signal to markets that the central bank anticipates a larger than expected contraction in the future. In their paper, Jarociński and Karadi (2020) make an effort to separately identify this information shock content of the ECB's monetary policy announcements in addition to the pure monetary policy component. In this section, we aim to show that our conclusions are robust to using their measure of pure monetary policy surprises as the shock variable.

Importantly, when estimating the effect of monetary policy on output, Jarociński and Karadi (2020) use their measure directly and do not use it to instrument for the surprise content in interest changes. We follow the same approach. Thus, in Equation 2, we substitute for $\Delta i$ and estimate the following regression using OLS:

$$\Delta \log(\overline{earn}^q_{t+h}) = \alpha_h + \beta^q_h S_{JK} + \theta X_t + \epsilon^q_{t+h} \tag{25}$$

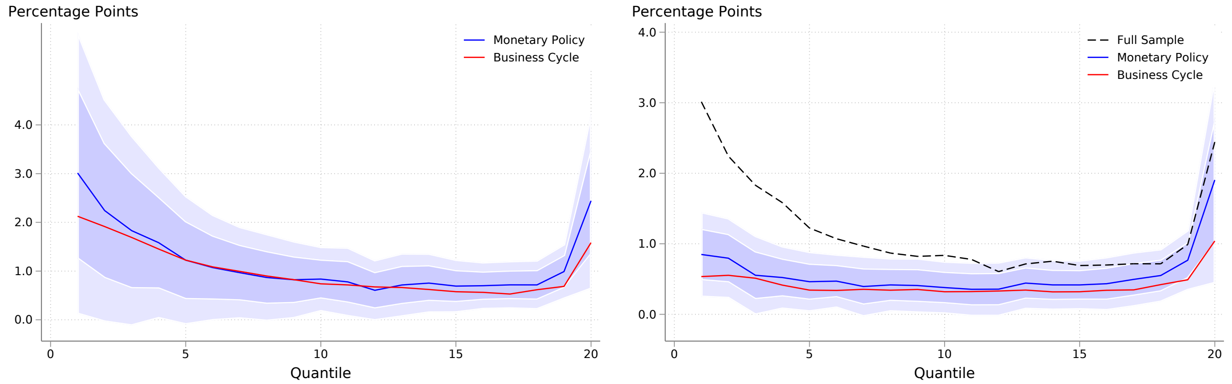
where $S_{JK}$ is the monetary policy shock measure from Jarociński and Karadi (2020). For the estimation, we use the same control variables as in our main specification: three lags of $\Delta i_t$, aggregate earnings and $S_{JK}$ as well as calendar month dummies.

Figure 13 reports the results. As in our main specification, monetary policy shocks have a larger effect at the low end of the aggregate income distribution. In fact, using the pure monetary policy component isolated in Jarociński and Karadi (2020), we find slightly larger effects on earnings, compared to the baseline. Again, the heterogeneity across the distribution is entirely driven by extensive margin transitions: the right panel of Figure 13 shows again homogeneous responses for employed individuals.

From this exercise we conclude that our results are robust to different identification schemes for monetary policy shocks.[22]

[22] We have also estimated regression 2 using the monetary policy surprises in Altavilla et al. (2019) and the results, not shown here, are very similar.

Figure 13: Regression coefficients $\beta^q_{12}$ across the income distribution




(a) Full sample

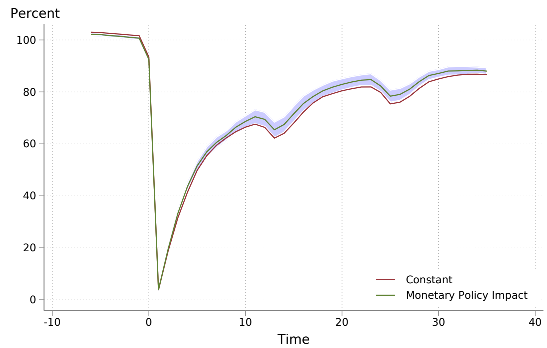
(b) Only individuals employed in $t-1$ and $t+12$

**Note:** The *Left Panel* plots the coefficients $\beta^q_{12}$ in Equation (2) (scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings) and $\beta^q_{Y,12}$ in Equation (3), separately for individuals who shared the same ventile of the permanent income distribution in period $t-1$. Income growth is computed as the log-change in the average income of individuals who were in the same ventile at time $t-1$. **For this exercise, we substitute our measure of monetary policy surprises with the one proposed by Jarociński and Karadi (2020).** The *Right Panel* compares the coefficients $\beta^q_{12}$ for the full sample in a ventile (gray dashed line) to $\beta^{q,E}_{12}$ and $\beta^{q,E}_{Y,12}$, estimated on a smaller sample of individuals who are employed both in period $t-1$ and $t+12$ (the blue and red lines, respectively). Ventiles are constructed based on average earnings during the five years prior to $t-1$, conditional on gender and five-year age brackets. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC robust standard errors. The sample period is 2000-2013.

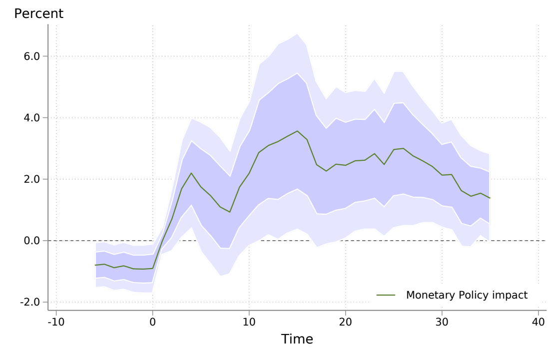
## B.3 Long-run effects of monetary policy on earnings

In section 4.4, we quantify the long-run impact of monetary policy on employment outcomes. Here, we conduct the same exercise, but for earnings (including zeros). Thus, we estimate Equation 6, using the average earnings by tercile. Figure 14 shows the results. The results in the top left panel represent earnings of individuals who become unemployed in period 0 relative to those who do not. Naturally, earnings approach zero in the first month after the transition into unemployment. However, some individuals find new jobs in the same month, implying positive average earnings. An accommodative monetary policy surprise steepens the slope of this recovery. As shown in the top right panel, such a shock increases earnings by about three percent, 20 months after unemployment. These effects are stronger than for the other two terciles, plotted in the bottom panel of the same figure. For the second decile, especially, the positive effect of monetary policy on earnings only materializes after close to two years.

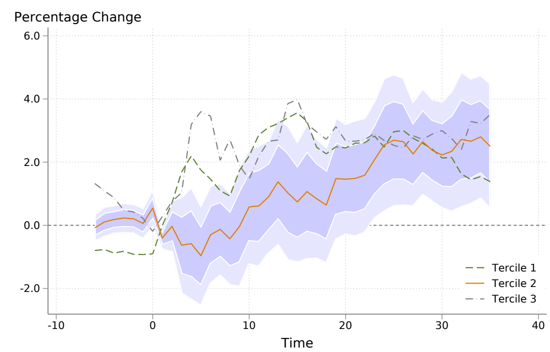
Figure 14: Effect of monetary policy on average earnings after unemployment




(a) First tercile




(b) First tercile – Difference



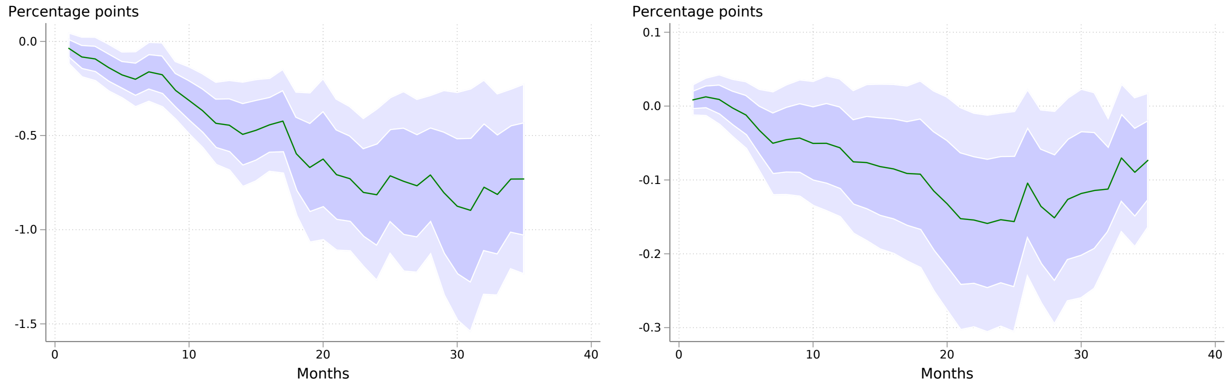

(c) Earnings

**Note:** The *Top Left Panel* shows the average earnings of individuals in the first tercile who transition into unemployment in month $t = 0$, relative to those who don't (red line). The green line shows how relative earnings change after a monetary policy surprise (scaled to be consistent with a one-percent increase in aggregate earnings). The shaded areas indicate 90 percent confidence intervals based on HAC standard errors. The *Top Right Panel* isolates the effect of a monetary policy surprise on the relative earnings, i.e. the difference between the two lines. The *Bottom Panel* shows the same effect for all three terciles. The dashed green lines represent estimates for the first tercile, the solid yellow lines those for the second and the dash-dotted grey line those for the third. The shaded areas in the top right and bottom graph represent 68 and 90 percent confidence bands based on HAC standard errors for the second tercile. Terciles are constructed based on average earnings during the five years prior to the start of the unemployment episode in period 0, conditional on gender and five-year age brackets. The sample period is 2000-2013.

## B.4 Impulse responses of sample aggregates

Figure 15 plots the impulse responses for aggregate earnings and the aggregate employment rate (as a percentage of fully attached individuals) as implied by our sample. The left panel plots the coefficient $\beta^h$ in Equation 2 where the variable on the left-hand side is average earnings across all individuals, by month. In response to a one-standard-deviation monetary policy surprise, aggregate earnings fall by 0.2 percent by the end of the first year after

Figure 15: Aggregate impulse responses




(a) Regression coefficients $\beta_h$ for the full sample (b) Regression coefficients $\gamma_h$ for the full sample

**Note:** The *Left Panel* plots the coefficients $\beta_h$ in Equation (2) for aggregate earnings in the full sample, scaled by a one-standard-error contractionary monetary policy surprise. The *Right Panel* plots the coefficients $\beta_h$ for the average employment rate in the sample. The shaded areas represent 68 and 90 percent confidence bands based on HAC standard errors, The sample period is 2000-2013.

the shock. The right panel plots the same estimation for the average employment rate. In response to a one-standard-deviation surprise, average employment falls by about 0.2 percentage points by the end of the first year after the shock.

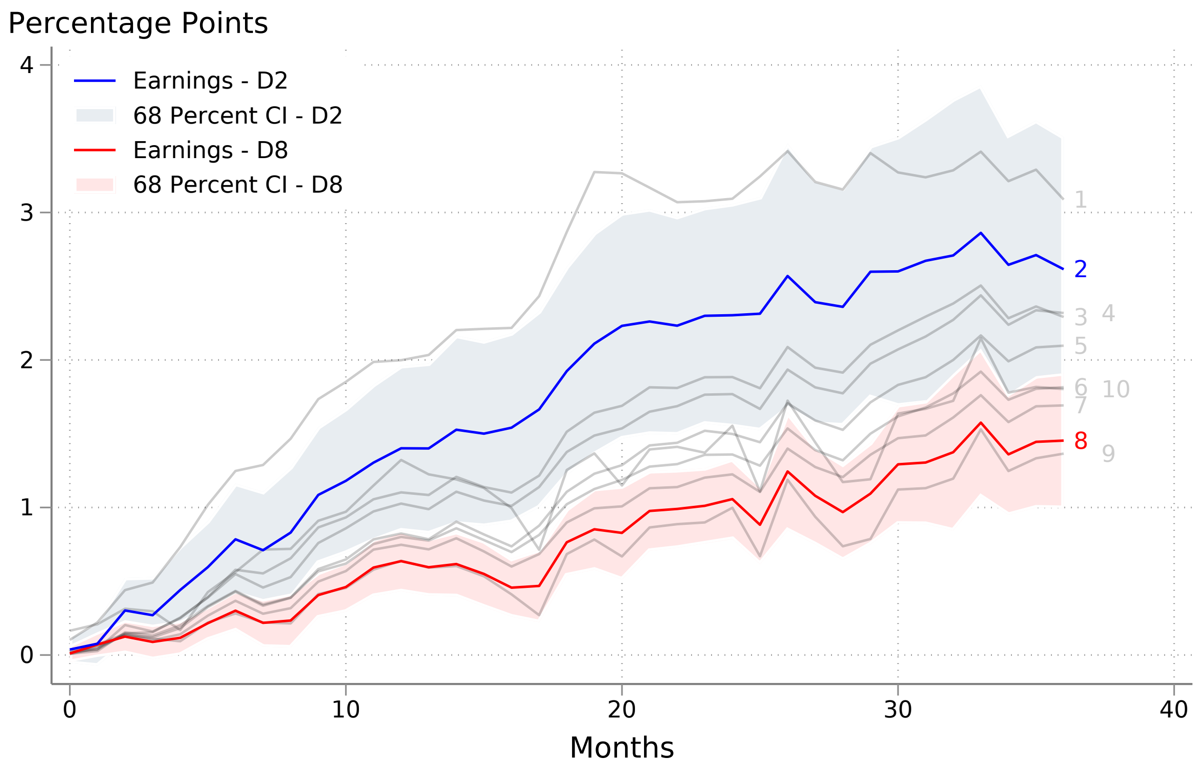
Both figures are in line with the aggregate responses reported in Figure 1: a surprising contraction in monetary policy depresses economic activity and decreases aggregate earnings and aggregate employment.

## B.5 Impulse responses by decile

In addition to the aggregate impulse responses, Figure 16 reports the earnings impulse responses by quantile. In order to not clutter the figure too much, we conduct this exercise using deciles, as opposed to the ventiles in the main analysis. Moreover, we only report confidence intervals for deciles two and eight. These impulse responses are also highlighted in blue and red, respectively. All other impulse responses are displayed in grey.

The figure shows that the pattern observed at the 12-month horizon is consistent even at longer horizons: at the bottom of the permanent income distribution, monetary policy has a stronger effect on earnings growth compared to the top. In fact, the difference increases over time, such that two years after the shock, earnings growth is about 0.2 percentage points lower for the second decile, compared to the eighth.

Figure 16: Regression coefficients $\beta_h^d$ by decile $d$

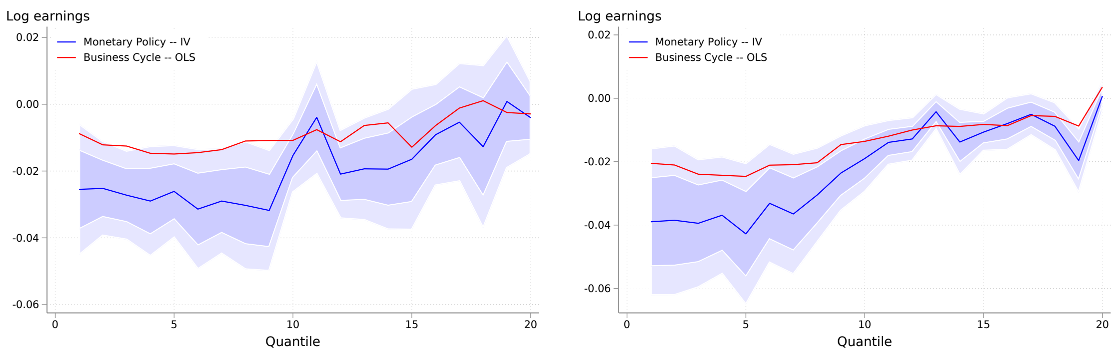



**Note:** The figure plots the coefficients $\beta_h^d$ in Equation (2), in response to a one-standard-error contractionary monetary policy surprise. Here, $d$ represents deciles, as opposed to the 20 quantiles $q$ in the main analysis. Each line represents the impulse response of a different decile, with the second decile highlighted in blue and the eighth decile highlighted in red. The shaded areas represent 68% confidence intervals.

## B.6 Earnings of job leavers and starters

In this section, we supplement the evidence in Section 4.1 by showing the effect of monetary policy on two remaining components of average earnings growth: exit-earnings of separators $(t-1)$ and starting-earnings job-finders $(t+12)$.

Figure 17: Components of the extensive margin




(a) Last earnings of separators

(b) First earnings of job-finders

**Note:** The *Left Panel* plots $\beta_{t-1}^{q,E\to U}$ from Equation 26 with $\Delta i$ as the right-hand side variable, instrumented using our monetary policy instrument (blue) and using $\Delta Y$ as the right-hand side variable (red). The *Right Panel* plots $\beta_{t+12}^{q,U\to E}$ from Equation 26 with $\Delta i$ as the right-hand side variable, instrumented using our monetary policy instrument (blue) and using $\Delta Y$ as the right-hand side variable (red). Ventiles are constructed based on average earnings during the five years prior to $t-1$, conditional on gender and five-year age brackets. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC robust standard errors. The sample period is 2000-2013.

To quantify the effect of a monetary policy surprise on the earnings of separators or job-finders, we estimate the following specification:

$$\log(\overline{earn}_{t+12,t-1}^{q,\tau}) = \alpha_h + \beta_h^{q,\tau}\Delta i_t + \theta X_t + \epsilon_{t+h}^q \tag{26}$$

where $\overline{earn}_{t+12,t-1}^{q,\tau}$ represents the average earnings of all individuals in quantile $q$ experiencing labor market transition $\tau \in \{E \to U; U \to E\}$ between periods $t$ and $t+12$. As before, $\Delta i$ represents changes in the ECB's policy rate which we instrument for as described in section 4.1; and $X_t$ is the vector of controls variables.

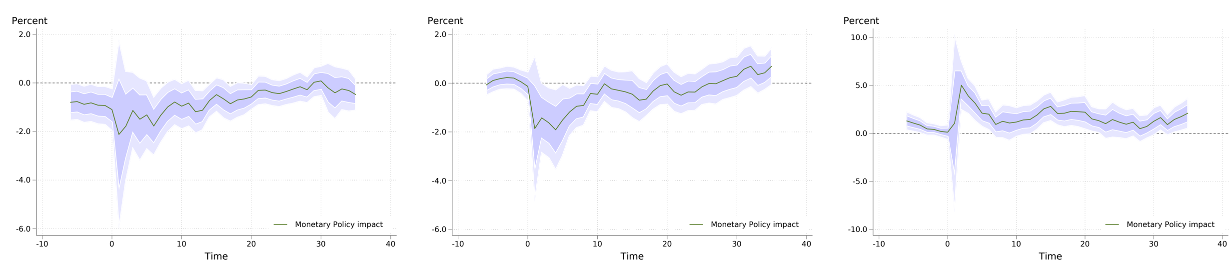
The bottom left panel of Figure 17 displays the log of the average earnings observation in period $t-1$ for individuals who transition from employment to unemployment by period $t+12$. The negative estimates for $\beta^{q,E\to U}$ imply that those individuals who separate during the year after an expansionary monetary policy shock have lower earnings (within ventile) than those who separate in other periods. The effect is strongest at the bottom of the distribution. This result is in line with the top left panel of Figure 3, which indicates that separation rates (the complement to the E-to-E rates reported in the Figure) fall most at the low end of the income distribution. Hence, those who separate after an expansionary shock are likely in the low tail of the earnings (or productivity) distribution.

The bottom right panel of Figure 17 reports the change in the log of average entry-earnings due to a monetary expansion, for those who transition from unemployment to employment. Entry-earnings appear to be dampened, implying that although job-finding rates increase homogeneously across the distribution (according to the top right panel in Figure 3), the bottom of the distribution appears to sacrifice earnings.

To foster a better understanding of these results, we take a longer-run approach and investigate how the earnings of individuals who become unemployed in period 0 evolve relative to the earnings of those who do not. This exercise is similar in spirit to the one conducted in section 4.4, however, here we condition on employment.

Figure 18 documents that, conditional on employment, the earnings difference between individuals who become unemployed in period 0 and those who do not is larger after an expansionary monetary policy regime. Still, Figure 18 documents that an expansion has positive effects on re-employment probabilities for those who become unemployed in period 0 (left panel), and that these appear to dominate the conditional earnings effects documented below (middle panel)

Figure 18: Conditional earnings effects of unemployment




(a) First tercile

(b) Second tercile

(c) Third tercile

**Note:** The figure shows the effect of a monetary policy surprise (scaled to be consistent with a one-percent increase in aggregate earnings), on the earnings of individuals who become unemployed in period 0 across terciles, relative to those who do not, **conditional on being employed**. The shaded areas represent 68 and 90 percent confidence bands based on HAC standard errors for the second tercile. Terciles are constructed based on average earnings during the five years prior to the start of the unemployment episode in period 0, conditional on gender and five-year age brackets. The sample period is 2000-2013. The *Left Panel* shows the effect in the first tercile. The *Middle Panel* shows the effect in the first tercile. The *Right Panel* shows the effect in the first tercile.

## B.7 Significance of differences across quantiles

From Figure 2, it is difficult to ascertain whether the impact of monetary policy is significantly different across quantiles. In this section, we report the t-statistics for such a test across all combinations of quantiles.

Table 10 reports the results for a t-test that investigates whether two quantile-specific coefficients are equal, i.e., the difference between them is zero. As the t-statistic grows, so does the likelihood that the null hypothesis is rejected.

The most significant differences can be observed between the first quantile and those above the median of the income distribution. As we move up the income distribution, the hypothesis of equality between two coefficients is less likely to be rejected; e.g., the coefficient reported for the third quantile is significantly different from coefficients for all quantiles between the 14th and 18th at the 68% significance level. At the same time, coefficients reported for the 16th quantile are singificantly different from those reported for the first seven quantiles as well as the last, at a significance level of 68%.

Table 10: Test for difference of quantile-specific coefficients

| | Q1 | Q2 | Q3 | Q4 | Q5 | Q6 | Q7 | Q8 | Q9 | Q10 | Q11 | Q12 | Q13 | Q14 | Q15 | Q16 | Q17 | Q18 | Q19 | Q20 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 2 | -.6 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 3 | -.8 | -.19 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 4 | -.99 | -.37 | -.19 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 5 | -1.19 | -.59 | -.43 | -.25 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 6 | -1.35 | -.76 | -.61 | -.45 | -.21 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 7 | -1.3 | -.69 | -.54 | -.36 | -.1 | .13 | 0 | . | . | . | . | . | . | . | . | . | . | . | . | . |
| 8 | -1.49 | -.91 | -.79 | -.65 | -.41 | -.21 | -.37 | 0 | . | . | . | . | . | . | . | . | . | . | . | . |
| 9 | -1.63 | -1.08 | -.98 | -.87 | -.66 | -.48 | -.66 | -.3 | 0 | . | . | . | . | . | . | . | . | . | . | . |
| 10 | -1.65 | -1.11 | -1.02 | -.91 | -.7 | -.53 | -.72 | -.35 | -.05 | 0 | . | . | . | . | . | . | . | . | . | . |
| 11 | -1.6 | -1.05 | -.95 | -.83 | -.61 | -.42 | -.6 | -.24 | .06 | .12 | 0 | . | . | . | . | . | . | . | . | . |
| 12 | -1.73 | -1.2 | -1.13 | -1.04 | -.85 | -.69 | -.9 | -.54 | -.24 | -.19 | -.31 | 0 | . | . | . | . | . | . | . | . |
| 13 | -1.75 | -1.22 | -1.15 | -1.07 | -.88 | -.73 | -.94 | -.58 | -.29 | -.24 | -.36 | -.06 | 0 | . | . | . | . | . | . | . |
| 14 | -1.72 | -1.19 | -1.11 | -1.02 | -.83 | -.67 | -.87 | -.51 | -.21 | -.16 | -.28 | .03 | .09 | 0 | . | . | . | . | . | . |
| 15 | -1.84 | -1.32 | -1.27 | -1.21 | -1.05 | -.91 | -1.14 | -.8 | -.51 | -.47 | -.58 | -.29 | -.22 | -.32 | 0 | . | . | . | . | . |
| 16 | -1.9 | -1.39 | -1.36 | -1.31 | -1.16 | -1.04 | -1.29 | -.96 | -.68 | -.64 | -.74 | -.47 | -.39 | -.5 | -.18 | 0 | . | . | . | . |
| 17 | -1.89 | -1.38 | -1.35 | -1.3 | -1.15 | -1.02 | -1.27 | -.94 | -.65 | -.61 | -.72 | -.45 | -.36 | -.47 | -.15 | .03 | 0 | . | . | . |
| 18 | -1.84 | -1.32 | -1.27 | -1.21 | -1.05 | -.91 | -1.14 | -.8 | -.52 | -.48 | -.58 | -.31 | -.24 | -.33 | -.02 | .15 | .12 | 0 | . | . |
| 19 | -1.66 | -1.13 | -1.04 | -.93 | -.73 | -.57 | -.74 | -.41 | -.15 | -.1 | -.2 | .06 | .11 | .04 | .3 | .44 | .42 | .31 | 0 | . |
| 20 | -.4 | .12 | .28 | .43 | .6 | .72 | .67 | .83 | .95 | .97 | .93 | 1.04 | 1.06 | 1.03 | 1.13 | 1.18 | 1.17 | 1.13 | .99 | 0 |

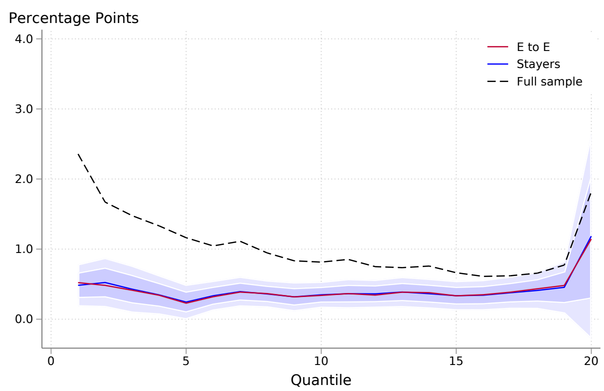
**Note:** The table reports the t-statistics of t-tests between pairs of quantiles. Each row-column combination represent pair. The diagonal is set to zero, the upper triangular is left blank to avoid repetition, as the test is symmetric.

## B.8 The impact of stayers

In Figure 2, to quantify the impact of the extensive margin on the effects of monetary policy, we restrict the sample to individuals who are employed in periods $t-1$ and $t+12$. This analysis conflates the effects of monetary policy on the earnings of job-switchers (those who are employed in $t$ and $t+1$, but not continuously at the same employer) and job-stayers (those with the same employer throughout the time between $t-1$ and $t+12$).

Figure 19 shows that this distinction is not meaningful in our context. The red line represents the effect of a monetary policy surprise on individuals who are employed in $t-1$ and $t+12$, the blue line restricts the sample further, to individuals who are job-stayers. The two lines are very close together, insignificantly different, and almost indistinguishable. The reason is that, in each quantile, the vast majority of individuals are job-stayers, conditional on being employed in $t-1$ and $t+12$. Thus, the earnings impact of job-switchers is low. Additionally, although more noisy, the earnings changes of job-switchers are similar to those of job-stayers (not shown).

Figure 19: The impact of switchers on coefficients $\beta_{12}^{q}$

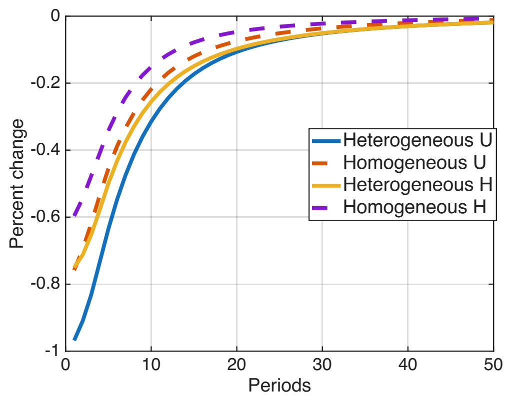



**Note:** The Figure plots the coefficients $\beta_{12}^{q}$ in Equation (2) (scaled by an expansionary monetary policy surprise consistent with a one-percent increase in aggregate earnings), separately for individuals who shared the same ventile of the permanent income distribution in period $t-1$. It compares the coefficients $\beta_{12}^{q}$ for the full sample across ventiles (gray dashed line) to $\beta_{12}^{q,E}$ and $\beta_{12}^{q,stay}$, estimated on a smaller sample of individuals who are employed both in period $t-1$ and $t+12$ and those who stay with the same employer for the same time (the blue and red lines, respectively). Ventiles are constructed based on average earnings during the five years prior to $t-1$, conditional on gender and five-year age brackets. The shaded areas indicate 68 and 90 percent confidence intervals based on HAC robust standard errors. The sample period is 2000-2013.

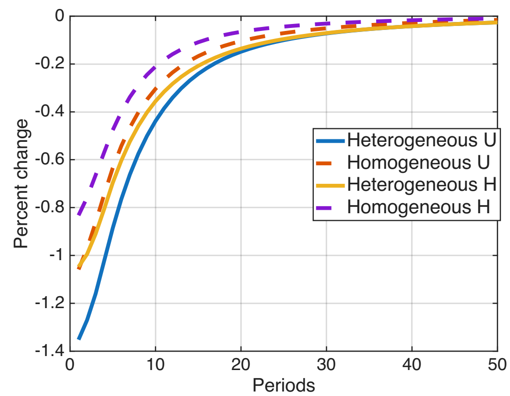
# C Full Model IRFs

Figure 20 shows the impulse responses of macroeconomic aggregates in the model to a 25 basis point monetary policy shock. The top left graph summarizes the output response for our four counterfactuals. The economy with heterogeneous incidence of unemployment ("Heterogeneous U", blue) responds strongest. Upon impact, output falls by close to one percent, before returning back to steady state. If the incidence of unemployment is homogeneous ("Homogeneous U", red-dashed), the initial rise in output is about 20% smaller. The patterns are the same for consumption. The reason for the large difference is the fact that with homogeneous incidence, the decrease in unemployment is more concentrated among highly productive workers. These agents have lower MPCs in both economies, but discount factors are lower in the heterogeneous unemployment case. Due to this, agents in the heterogeneous risk economy adjust their consumption more in response to decreases in employment risk in general. Beyond this, the decrease

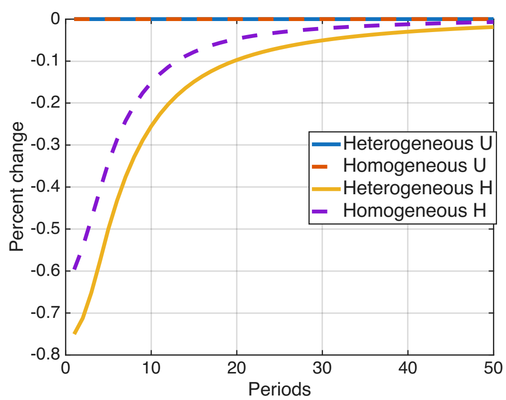
Figure 20: General equilibrium impulse responses



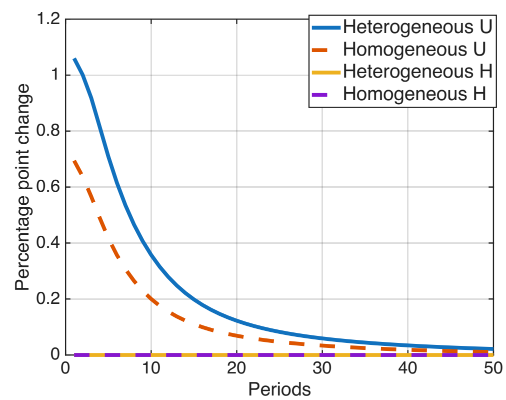

(a) Output



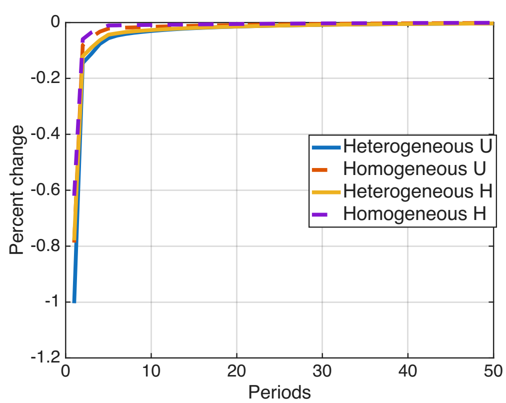

(b) Consumption



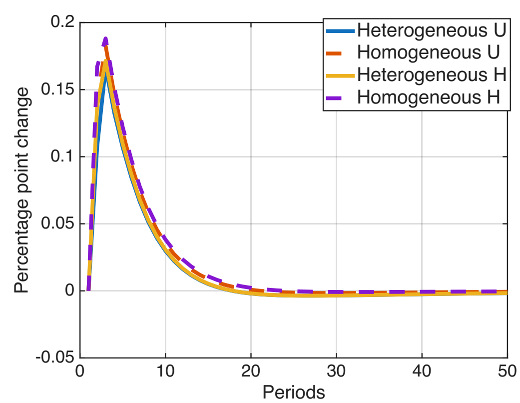

(c) Hours




(d) Unemployment




(e) Real wage




(f) Interest Rate

**Note:** The figure shows the model generated impulse responses to a 25 basis point monetary policy contraction for four economies: the benchmark economy (Heterogeneous U, solid blue), an economy with equal incidence of unemployment in steady state and along the transition (Homogeneous U, dashed red), an economy with unequal unemployment incidence in steady state and hours adjustments along the transition (Heterogeneous H, solid yellow), and an economy with equal unemployment incidence in steady state, but hours adjustments along the transition (Homogeneous H, dashed purple).