The Regional Keynesian Cross†

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March 2025

Abstract

We study how regional heterogeneity shapes the aggregate transmission of monetary policy and its distributional implications across space. We build a multi-region Heterogeneous-Agent New Keynesian model with 3,140 U.S. counties and cross-county differences in (i) intertemporal Marginal Propensities to Consume (MPCs) and (ii) non-tradable employment shares. We analytically characterize the nationwide consumption response to monetary policy in terms of the joint distribution of (i) and (ii). Using U.S. and Italian micro-data, we construct novel empirical measures of regional MPCs to validate our theory. Quantitatively, geographic heterogeneity leads to large distributional consequences of monetary policy across space and can sizably amplify its aggregate effects.

JEL Codes: E12, E21, E23, E52, F41.

Keywords: Monetary policy, Keynesian multiplier, trade openness, monetary union.

†We thank Klaus Adam, Adrien Auclert, Bence Bardóczy, Christian Bayer, Adrien Bilal, Ambrogio Cesa-Bianchi, Luis Calderon (discussant), Jeff Campbell (discussant), Edouard Challe (discussant), Gabriel Chodorow-Reich, Giancarlo Corsetti, Leonardo D'Amico, Maarten De Ridder, Wouter Den Haan, Matthias Doepke, Jan Eeckhout, Axelle Ferriere, Jordi Galí, Manuel García-Santana, Joe Hazell, Erik Hurst, Tomer Ifergane, Ethan Ilzetzki, Guido Lorenzoni, Greg Mankiw, Michael McMahon, Isabelle Mejean, Kurt Mitman, Ben Moll, Tommaso Monacelli, Gernot Müller, Dima Mukhin, Ricardo Reis, Nicolò Rizzotti, Kathrin Schlafmann (discussant), Adi Soenarjo, Ludwig Straub, Gianluca Violante, Iván Werning and seminar participants at various venues for useful comments and suggestions. Oliver Seager provided excellent research assistance. Bellifemine and Couturier acknowledge support from the STICERD Hayek grant. Jamilov acknowledges support from the Wheeler Institute for Business and Development and the John Fell Fund (CUD00490). Regional MPC data for U.S. and Italy are available at: www.regionalkeynesiancross.com. First Draft: December, 2022.

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1 Introduction

How does monetary policy impact aggregate fluctuations? Textbook macroeconomic models assume that monetary policy has uniform effects across space. However, extensive empirical work has shown that this is not the case.[1] Large geographical disparities are a defining characteristic of many modern economies, causing different local areas to respond to the same aggregate shock in markedly different ways. This implies that policies designed to stabilize the national economy may also affect regional inequalities. It also suggests that the geography of local economic characteristics may be important for understanding the aggregate effects of monetary policy.

In this paper, we study how regional heterogeneity affects monetary policy transmission in a multi-region model of the U.S. economy. Our framework features 3,140 counties and regional asymmetry along two measurable dimensions: intertemporal Marginal Propensities to Consume (iMPCs) and openness to trade. We show that the joint distribution of regional iMPCs and trade openness shapes monetary policy transmission both in the aggregate and across space. At the regional level, a Keynesian multiplier operates via the non-tradable labor market and is determined not only by iMPCs but also by the relative size of the non-tradable sector itself. Hence, our theory explicitly predicts systematic differences in the response to monetary policy across regions.

Moreover, because of the non-linear nature of the regional multiplier, we show analytically that local iMPCs and trade openness jointly determine the *nationwide* consumption response to monetary policy shocks. Finally, we construct a novel measure of regional MPCs for the U.S. and Italy in order to quantify and validate our model. We show that our calibrated model is quantitatively consistent with the data and can generate substantive aggregate amplification of monetary shocks.

Theory. Our model builds on two foundational concepts. The first is the dynamic Keynesian multiplier, which controls the equilibrium feedback loop between consumption and income changes and is a central mechanism of monetary policy transmission in the Heterogeneous-Agent New Keynesian (HANK) framework. The second is non-tradability—a standard source of geographical differentiation in international macroeconomics and the optimum currency area literature.

We merge these two insights in a multi-region HANK model of a monetary union that is populated by a continuum of counties. There are two types of goods: a tradable good that is traded nationally and a non-tradable good which has to be produced locally.

[1] See, for example, Carlino and Defina (1998) for an early analysis of the U.S. and Almgren et al. (2022) for the case of the euro area.

Counties are linked via the tradable goods market as well as a national bonds market. There is rich heterogeneity both within and across counties. Inside a county, households face uninsured idiosyncratic income risk. This gives rise to a distribution of income and wealth. Across counties, regions differ both in their local iMPCs and in the share of labor income in the non-tradable sector. Importantly, we go beyond studying individual counties in isolation and allow national demand to be determined endogenously by the continuum of asymmetric regions.

The regional Keynesian cross. We start our analysis by deriving closed-form expressions in the sequence space for the county-level employment and consumption responses to a monetary policy shock: the regional Keynesian cross. This is the currency union analogue of the intertemporal Keynesian cross in the closed-economy model of Auclert et al. (2024b) and its small and large open-economy counterparts in Auclert et al. (2021b) and Aggarwal et al. (2023), respectively. Our analytical expressions provide a clear decomposition of the total county-level response into just two channels: national equilibrium effects and regional equilibrium effects. The national equilibrium channel consists of the local response that is due to the change in national demand for tradables induced by the monetary shock. The regional equilibrium effect, in turn, consists of an intertemporal substitution channel—which captures the direct consumption response of local households to the interest rate change—and a regional Keynesian multiplier that summarizes the indirect effects via labor income. By governing the county-level exposure to local versus national demand, the share of non-tradable employment controls the relative importance of these two channels. As a result, the multiplier term is shaped not only by iMPCs but also by the non-tradable share of labor income.

Crucially, iMPCs and the non-tradable share enter *multiplicatively* in the expression for the regional Keynesian multiplier. This reflects the idea that the extent to which demand spills back to local labor markets depends jointly on (i) the pass-through of labor income to consumption, captured by the iMPCs, and (ii) the exposure of local labor income to local consumption, driven by the non-tradable share. Thus, spatial heterogeneity captured by these two variables endogenously translates into differences in county-level responses to the same monetary impulse.

The national Keynesian cross. Our main theoretical contribution is to move beyond county-level analysis to study the implications of regional heterogeneity for the aggregate effects of monetary policy. As the regional Keynesian multiplier is non-linear in iMPCs and the non-tradable share, regional heterogeneity does not wash out in the aggregate.

To show this, we integrate the continuum of regional Keynesian crosses over all counties. This yields the national Keynesian cross: an analytical characterization of the national employment and consumption responses in terms of the joint regional distribution of iMPCs and trade openness. The key statistics entering these expressions are cross-regional covariances between (transformations of) iMPCs and the share of non-tradable labor income.

Our national Keynesian cross result provides an intuitive decomposition of the aggregate consumption response into a direct and an indirect channel, just like in the standard intertemporal Keynesian cross. Regional heterogeneity enters this expression by modifying the magnitudes of both of these channels. Thus, the national Keynesian cross is a useful framework for studying how regional heterogeneity affects the aggregate response to monetary policy, both in terms of its magnitude and its underlying transmission mechanism. A key implication of this result is that two economies that share the same average nationwide iMPC may have very different responses to the same aggregate shock if they feature different regional distributions.

Empirical evidence. An important advantage of our theory is that the key statistics governing the extent to which regional heterogeneity affects aggregate fluctuations can be measured directly in the data. We now turn to detailed county-level U.S. data in order to measure these statistics and to empirically evaluate key mechanisms of our model. First, we develop a novel methodology to estimate county-level MPCs. Our approach is related to Patterson (2023) and leverages self-reported MPCs from Fuster et al. (2020) together with regional data from the American Community Survey to back out a synthetic measure of MPCs at the county level.[2] We also construct county-level measures of the share of non-tradable employment and merge them with our regional MPC data.

Second, in line with the state-level evidence from Carlino and Defina (1998), we document that identified monetary policy shocks induce local employment responses that vary substantially across U.S. counties. Third, this variation is not just noise as it can be systematically explained by our two regional characteristics. In particular, we find that counties that have high MPCs are more responsive to monetary shocks, as was also shown by Almgren et al. (2022) in the context of the euro area. We also find that counties with a high non-tradable share of employment are more responsive to monetary shocks. Furthermore, the slope of the relationship between MPCs and the local employment response is increasing in the non-tradable share, which is consistent with our theory.

[2]Our approach is also related to recent advancements in the measurement of MPCs using large-scale survey data (Colarieti et al., 2024) and randomized experiments (Boehm et al., 2025).

Quantitative implications. Finally, we bring our model to the data in order to assess its quantitative implications. We develop a novel computational approach that enables us to calibrate our framework to the full set of 3,140 U.S. counties. This constitutes one of the contributions of our paper. We then validate our theory by showing that it is able to match our empirical findings. Specifically, we replicate our empirical regressions using model-generated data and find that counties with either higher local MPCs or a larger share of non-tradable employment show greater responsiveness to monetary policy. Moreover, we find the interaction term between these two dimensions to be economically significant.

We also leverage our regional Keynesian cross result to demonstrate that the monetary transmission mechanism itself varies significantly across space. We illustrate this through three model-based case studies of real-world U.S. counties: New York County (New York), Luce County (Michigan), and Putnam County (West Virginia). We select these counties for their starkly contrasting household and industrial compositions: New York County is characterized by very low MPCs and a high share of non-tradables, Putnam County by very high MPCs and a low share of non-tradables, and Luce County by both high MPCs and a high share of non-tradables. Consequently, these counties not only respond differently to monetary policy shocks but do so through distinct channels which we elicit.

Finally, our theory predicts that distinct regional distributions can induce different nationwide implications of the same aggregate shock. We demonstrate this insight quantitatively by calibrating our model to two different geographies: the U.S. and Italy. In doing so, we compute a measure of regional MPCs using detailed Italian survey data. Our findings reveal that regional heterogeneity has a limited effect on the aggregate response to monetary policy in the U.S. In contrast, properly accounting for the regional distribution in Italy amplifies the aggregate response to monetary shocks by as much as 18% on impact and 59% cumulatively. These results highlight how patterns of spatial heterogeneity shape the extent to which regional distributions influence aggregate dynamics. Consequently, policymakers should consider and monitor geographic distributions of household and industry compositions to better evaluate the overall effectiveness of their monetary policy instruments.

Related literature. We are building on the burgeoning HANK literature that develops micro-founded, non-Ricardian environments with household heterogeneity and nominal rigidities (McKay and Reis, 2016, Kaplan et al., 2018, Auclert et al., 2024b).[3] In particular, our study is most closely related to those focusing on monetary policy, which is the primary interest of our paper (McKay et al., 2016, Kaplan et al., 2018, Auclert, 2019, Ravn and

[3]See Auclert et al. (2024a) for a recent review of this rapidly growing literature.

Sterk, 2020, Luetticke, 2021, Kekre and Lenel, 2022, Wolf, 2025). The seminal neutrality result of Werning (2015) states that HANK models can, in the aggregate, generally behave as if they were governed by a representative agent. A critical result of our paper is to show quantitatively that this is not the case in our model—regional heterogeneity can sizably amplify the effects of monetary shocks due to the non-linearity of the regional multiplier.[4]

In parallel with the closed-economy HANK literature, a growing body of work explores open-economy HANK models following the seminal Galí and Monacelli (2005) small-open-economy approach (de Ferra et al., 2020, Auclert et al., 2021b, Guo et al., 2023).[5] For example, de Ferra et al. (2020) develop a HANK framework of a small open economy with household heterogeneity to study the transmission of foreign shocks such as sudden stops in capital inflows. In this line of work, the analysis generally centers around an individual region while the national (or global) demand for goods and services is assumed to remain fixed. Our key contribution is to go beyond the study of a singular county in isolation and to consider a world where national demand is fully endogenous and determined by the entire asymmetric distribution of regional demands—which is our national Keynesian cross result.

An important exception to both the closed-economy and open-economy HANK literatures—and the paper most closely related to ours—is Aggarwal et al. (2023). Aggarwal et al. (2023) develop the first many-country model of large open economies. While complementary, our paper differs in several substantive ways. First and foremost, we analytically characterize the general-equilibrium national response to a monetary policy shock as a closed-form function of cross-regional asymmetries in iMPCs and trade openness. Second, whereas Aggarwal et al. (2023) study fiscal policy, we focus exclusively on monetary policy. Third, while Aggarwal et al. (2023) calibrate their model using aggregate data on home bias and fiscal policy for 26 countries, we calibrate our economy to 3,140 U.S. counties using county-level data on MPCs and non-tradable employment. Fourth and finally, our newly constructed MPC data for the U.S. and Italy represents an additional contribution of the paper. As such, to the best of our knowledge, we are the first to study monetary policy in a calibrated multi-region model with many asymmetric economies.

An alternative non-Ricardian approach to HANK is the Two-Agent New Keynesian (TANK) class of models (Galí et al., 2007, Bilbiie, 2008, 2020, Debortoli and Galí, 2024) and

[4]A common thread in the HANK literature has also been the emphasis on non-linearities in agents' *partial-equilibrium* policy functions to illustrate the role of distributions for aggregate dynamics (Kaplan et al., 2018, Ottonello and Winberry, 2020). We shift the focus to non-linearities arising from local *general-equilibrium* effects as the key driver of heterogeneity.

[5]See also Zhou (2022), Oskolkov (2023), Druedahl et al. (2024a,b), and Bellifemine et al. (2025).

the recently developed tractable-HANK framework (Bilbiie, 2024). We adopt the HANK modeling approach in our paper for two main reasons. First, incorporating market incompleteness and the full distribution of agents ensures stationarity properties that would likely not hold in the TANK special case of our setup (Ghironi, 2006). Second, quantitative HANK frameworks match empirical iMPCs much more accurately (Fagereng et al., 2021, Auclert et al., 2024b).

As discussed at the beginning of the paper, the first key object in our theory is the general-equilibrium multiplier channel, particularly in its dynamic form as in Bilbiie (2020) and Auclert et al. (2024b). Variants of this channel have been studied in, among others, Kaplan et al. (2018), Auerbach et al. (2020), Chodorow-Reich et al. (2021), Schaab and Tan (2023), Aggarwal et al. (2023), Auclert et al. (2024b), Wolf (2023, 2025). The second—openness to trade—is a classic source of geographic fragmentation in the optimum currency area literature (Mundell, 1961, McKinnon, 1963, Kenen, 1969). Relatedly, our model also connects with the theoretical literature studying monetary and fiscal unions (Farhi and Werning, 2016, 2017, Pica, 2023, Bayer et al., 2024, D'Amico and Alekseev, 2024, Boehnert et al., 2025).

At the core of our analysis is the basic argument that aggregate fluctuations are driven by underlying regional economic cycles (Beraja et al., 2018, 2019) and an asymmetric spatial distribution of economic activity (Bilal and Rossi-Hansberg, 2021, Bilal, 2023, Oberfield et al., 2024). Our theoretical treatment builds on the sufficient statistics approach, which is a method of characterizing complex general-equilibrium dynamics with a small set of measurable objects, and popularized by, among others, Kaplan and Violante (2014), Auclert (2019), Berger et al. (2021), Auclert and Mitman (2023). Relative to the above, our contribution is to embed the canonical Keynesian multiplier into a multi-regional framework with thousands of localities and to quantify it with just two sufficient statistics—regional iMPCs and non-tradable shares of labor income—which we measure in the data.

Our empirical analysis contributes to a large body of empirical work that studies how monetary policy operates across space within a currency union (Carlino and Defina, 1998, De Ridder and Pfajfar, 2017, Corsetti et al., 2021, Almgren et al., 2022, Adam et al., 2022, de Groot et al., 2023). We contribute to this literature in two main ways. First, we provide a novel county-level measure of MPCs for the U.S. Second, we document that (i) counties with a larger non-tradable sector are more responsive to monetary policy shocks and (ii) the gradient of the relationship between MPCs and the local employment response is increasing in the non-tradable share.

Finally, we cast our framework in the sequence space, an influential method of writing and solving models, developed by Mankiw and Reis (2007), Boppart et al. (2018), and

Auclert et al. (2021a).

Outline. The rest of the paper is organized as follows. In Section 2 we lay out our model and derive the main theoretical results on the regional Keynesian cross and the national Keynesian cross. Section 3 introduces our methodology to estimate local MPCs and non-tradable employment and presents our empirical findings. In Section 4, we bring the model to the data and describe our calibration strategy. Section 5 presents our quantitative results both for the distributional effects of monetary policy and for the aggregate implications of regional heterogeneity. Section 6 concludes.

2 Theory

In this section, we first introduce our multi-region Heterogeneous Agent New Keynesian (HANK) model of a monetary union. Next, we cast our modeling framework in the sequence space and define two objects that will be useful in our analysis: the sequence space Jacobian matrices capturing intertemporal Marginal Propensities to Consume (iMPCs) à la Auclert et al. (2024b) and the share of non-tradable labor income. Finally, we derive our two main theoretical results. First, the regional Keynesian cross, which characterizes how iMPCs and the share of non-tradable employment shape the heterogeneity in regional responses to monetary shocks. And second, the national Keynesian cross, describing how the joint distribution of iMPCs and non-tradable employment across space affects the aggregate, nationwide transmission of monetary policy.

2.1 Model environment

Time $t \geq 0$ is discrete. There is a continuum of atomistic counties (regions) indexed by $j \in [0,1]$ and modeled as small open economies à la Galí and Monacelli (2005). Counties belong to a monetary union. There is no aggregate uncertainty. For transitional dynamics, we consider perfect-foresight impulse responses to shocks around the steady state ("MIT shocks").

Households. Each county j is inhabited by a continuum of households $i \in [0,1]$. As in the standard incomplete markets model à la Bewley (1977) and Huggett (1993), households are ex-ante identical, but face non-insurable idiosyncratic shocks to their labor productivity e_{jit}, which evolves over time according to a Markov process. Preferences of household i living in county j are defined over an aggregate consumption good c_{jit} as well as aggregate labor supply ℓ_{jit}, which imply the following time-0 utility:

$$\mathbb{E}_0 \sum_{t \geq 0} \beta_j^t \{u(c_{jit}) - v(\ell_{jit})\} \tag{1}$$

where $\beta_j \in (0,1)$ is the discount factor. Agents can imperfectly insure themselves by trading in a nominal risk-free bond with real value b_{jit} subject to a borrowing limit $\underline{b}_j \leq 0$.

$$c_{jit} + b_{jit+1} = z_{jit} e_{jit} + \frac{1+i_t}{1+\pi_{jt}} b_{jit}, \quad b_{jit+1} \geq \underline{b}_j \tag{2}$$

In (2) above, $z_{jit} \equiv \frac{W_{jt}}{P_{jt}} \ell_{jit}$ is real gross labor income, where W_{jt} and P_{jt} are, respectively, the aggregate wage and price index in county j, both of which are taken as given and will be defined momentarily. Additionally, we denote by $\pi_{jt} \equiv \frac{P_{jt} - P_{jt-1}}{P_{jt-1}}$ price inflation in county j and by i_t the nominal interest rate on bonds.[6]

As can be seen from (1) and (2), we allow counties to differ in their degree of patience β_j, process for labor income productivity e_{ijt}, and credit market structure, as captured by the borrowing constraint $\underline{b}_j$. Cross-county differences in these three parameters are potential sources of regional iMPC heterogeneity in our framework.

Sectoral composition. There are two consumption goods in the economy: non-tradables and tradables. The two goods are combined into the aggregate consumption basket c_{jit} according to a constant-elasticity-of-substitution (CES) aggregator:

$$c_{jit} = \left[\omega_j^{1/\nu} \left(c_{jit}^{NT} \right)^{(\nu-1)/\nu} + (1-\omega_j)^{1/\nu} \left(c_{jit}^{T} \right)^{(\nu-1)/\nu} \right]^{\frac{\nu}{\nu-1}} \tag{3}$$

where c_{jit}^{NT} and c_{jit}^{T}, respectively, denote consumption of the non-tradable and tradable goods, ω_j is a county-specific parameter governing households' preferences for non-tradables, and $\nu > 0$ is the elasticity of substitution between the two types of goods. The defining feature of non-tradable goods is that they must be consumed in the same county where they have been produced. The tradable good is itself defined as a composite of different tradable varieties produced in each county j, as in Galí and Monacelli (2005, 2008):

$$c_{jit}^{T} = \left(\int_0^1 c_{jit}^{T}(j')^{\frac{\theta-1}{\theta}} dj' \right)^{\frac{\theta}{\theta-1}} \tag{4}$$

[6] As we explain further below, the nominal interest rate is equalized across counties because there exists a single nationally integrated asset market.

with $\theta > 0$ being the elasticity of substitution between tradable goods produced in different counties. This implies the following demand for tradables produced in county j' from residents of county j:

$$c_{jt}^{T}(j') = \left(\frac{P_t^T(j')}{P_t^T}\right)^{-\theta} c_{jt}^{T} \tag{5}$$

where $c_{jt}^{T} = \int c_{jit}^{T} di$ is aggregate tradable consumption from residents of county j, $P_t^T(j')$ is the price of the tradable variety produced in county j' and $P_t^T = \left(\int_0^1 P_t^T(j')^{1-\theta} dj'\right)^{\frac{1}{1-\theta}}$ is the price index for tradables. As is standard, households split their spending between the two types of goods as follows:

$$c_{jit}^{NT} = \omega_j \left(\frac{P_{jt}^{NT}}{P_{jt}}\right)^{-\nu} c_{jit} \quad \text{and} \quad c_{jit}^{T} = (1-\omega_j)\left(\frac{P_t^T}{P_{jt}}\right)^{-\nu} c_{jit} \tag{6}$$

where P_{jt}^{NT} is the price for non-tradables in county j and P_{jt} denotes county j's aggregate price index, which is given by:

$$P_{jt} = \left[\omega_j \left(P_{jt}^{NT}\right)^{1-\nu} + (1-\omega_j)\left(P_t^T\right)^{1-\nu}\right]^{\frac{1}{1-\nu}} \tag{7}$$

Because in our model preferences are homothetic and do not depend on the household type i, both the price and wage indices as well as the composition of the consumption basket are identical across households within any given county.

Similarly to demand, the supply side of each county j is composed of two sectors: one produces county j's tradable variety while the other produces the non-tradable good. We follow Berger et al. (2022) when modelling the supply of labor to the two sectors: individual households' aggregate labor supply ℓ_{jit} is a composite of a measure of labor supplied to the non-tradable sector, ℓ_{jit}^{NT}, and a measure ℓ_{jit}^{T}, which is supplied to the tradable sector. In particular, the labor supply composite is given by the following CES aggregator:

$$\ell_{jit} = \left(\alpha_j^{-\frac{1}{\eta}} (\ell_{jit}^{NT})^{\frac{\eta+1}{\eta}} + (1-\alpha_j)^{-\frac{1}{\eta}} (\ell_{jit}^{T})^{\frac{\eta+1}{\eta}}\right)^{\frac{\eta}{\eta+1}} \tag{8}$$

where $\eta > 0$ is the elasticity of labor substitution across sectors and is assumed to be constant across counties. The parameter α_j can be thought of as the endowment of workers in the non-tradable sector in county j and is allowed to vary across counties. Given (8), households allocate their labor as follows:

$$\ell^{NT}_{jit} = \alpha_j \left(\frac{W^{NT}_{jt}}{W_{jt}} \right)^{\eta} \ell_{jit}, \quad \text{and} \quad \ell^{T}_{jit} = (1-\alpha_j) \left(\frac{W^{T}_{jt}}{W_{jt}} \right)^{\eta} \ell_{jit} \tag{9}$$

where W^{NT}_{jt} and W^{T}_{jt} represent, respectively, the wage paid to workers in the non-tradable and tradable sector in county j and W_{jt} is the aggregate wage index, which is given by:

$$W_{jt} = \left[\alpha_j (W^{NT}_{jt})^{1+\eta} + (1-\alpha_j)(W^{T}_{jt})^{1+\eta} \right]^{\frac{1}{1+\eta}} \tag{10}$$

As is clear from (3) and (8), we allow counties to differ both in their labor endowment in the non-tradable sector—captured by α_j—as well as in the parameter ω_j, reflecting local consumers' preferences for non-tradables. These two objects will drive differences in the steady-state share of non-tradable employment across counties in our model.

Final good producers. Firms in both sectors produce using a linear production technology: $Y^s_{jt} = L^s_{jt}$, $s \in \{NT, T\}$. Moreover, in both sectors the market for final goods is perfectly competitive. As a result, final prices for the two goods equal marginal costs, i.e., $P^s_{jt} = W^s_{jt}$ and firms in all counties and sectors make zero profits in every period.

Labor markets. We introduce nominal rigidities in the form of sticky wages. In line with the New Keynesian sticky-wage literature (Erceg et al., 2000, Schmitt-Grohé and Uribe, 2005, Auclert et al., 2024b), we assume that the amount of hours worked by each household is determined by county-level labor unions. In every county j, there is a continuum of unions in each sector $s \in \{NT, T\}$ that choose wages and hours worked to maximize the welfare of the average household in that county. Unions allocate labor hours among their members in a uniform fashion, which implies that $\ell_{jit} = L_{jt}$.[7] We lay out the full union problem and derive the associated county-specific sectoral New Keynesian Phillips Curves in Appendix B.

Financial markets and monetary policy. There is a single asset traded in the economy: a nominally risk-free bond which is in zero net supply. Financial markets are fully integrated across the monetary union so that the nominal interest rate on the bond, i_t, is pinned down nationally and equalized across counties. Note that in our model, due to market incompleteness within each county, aggregate county-level asset holdings return

[7]The assumptions that the union maximizes the welfare of the average household as well as the uniform labor allocation rule can be easily relaxed to more general cases.

to steady state after a shock. As discussed by Ghironi (2006), with incomplete markets, the steady-state growth rate of consumption depends on aggregate asset holdings, which are uniquely determined in the steady state. In other words, the stationarity of our model is assured because market incompleteness results in an upward sloping asset supply schedule at the county level. This would not be the case in a two-agent New Keynesian (TANK) model, which is one reason why we use a HANK framework. Finally, there is a national central bank controlling the nominal interest rate i_t according to a standard Taylor rule.

2.2 Equilibrium

We now define the equilibrium concepts used in this paper. We distinguish between a regional equilibrium, which relies on a small-open-economy structure and takes national prices and quantities as given, and a national equilibrium, a situation in which *every* county is in a regional equilibrium and national markets clear. This distinction will prove useful for obtaining closed-form solutions and conducting a number of quantitative exercises, as shown later in the paper.

Steady states. We start by defining the steady-state equilibrium.

Definition 1 (Regional steady state)**.** *A steady-state regional equilibrium in county j is a set of quantities and prices $\{C_j, C_j^{NT}, C_j^T, L_j, L_j^{NT}, L_j^T, P_j, P_j^{NT}, P_j^T, W_j, W_j^{NT}, W_j^T\}$, a regional distribution over bonds and productivity $G_j(b,e)$, and a policy function $c_j(b,e)$, such that, given the interest rate on bonds r, national demand for the tradable good $\{C_{j'}^T\}_{j'}$, and the national tradable price P^T, households, unions, and firms in county j optimize, and regional markets clear:*

$$L_j^{NT} = C_j^{NT} \tag{11}$$

$$L_j^T = \int_0^1 \left(\frac{P_j^T}{P^T}\right)^{-\theta} C_{j'}^T dj' \tag{12}$$

Definition 2 (National steady state)**.** *A steady-state national equilibrium is a set of regional quantities and prices $\{C_j, C_j^{NT}, C_j^T, \{C_j^T(j')\}_{j'}, L_j, L_j^{NT}, L_j^T, P_j, P_j^{NT}, P_j^T, W_j, W_j^{NT}, W_j^T\}_j$, a national real interest rate on bonds r, a set of regional distributions over bonds and productivity $\{G_j(b,e)\}_j$, and policy functions $\{c_j(b,e), b_j'(b,e)\}_j$, such that all counties are in a regional equilibrium and national asset markets, and tradable goods markets clear:*

$$0 = \int_0^1 P_j \int_0^1 b_{ij} di \, dj \tag{13}$$

Transition dynamics. For analytical transparency and tractability, when analyzing equilibrium transitions following exogenous shocks, we assume throughout the main body of the paper that wages are rigid. In Appendix B, we relax the rigidity assumption, describe the equilibrium definition under sticky wages, and show that both our theoretical and quantitative results remain unaffected. Because wages are not allowed to adjust, we assume that output is demand-determined at the sector level. As a result, unions will be off their optimal labor supply equation and households will be off their optimal sectoral labor allocation condition. Moreover, under rigid prices the path for real interest rates is the same across counties and equal to that of the nominal interest rate, i.e., $r_{jt} = i_t$, since price inflation is equal to zero by definition.

Definition 3 (Regional equilibrium)**.** *Given an initial regional distribution $G_{j0}(b,e)$ over bonds b and idiosyncratic labor productivity e, an exogenous path of national demand for tradable goods $\{C_t^T\}_{t\geq 0}$, the steady-state price index for tradable goods P^T, and an exogenous path of real interest rates $\{r_t\}_{t\geq 0}$, a regional equilibrium for county j is a sequence of county j's aggregate quantities $\left\{L_{jt}, L_{jt}^{NT}, L_{jt}^T, C_{jt}, C_{jt}^{NT}, B_{jt}\right\}_{t\geq 0}$, individual allocation rules $\{c_{jt}(b,e), b_{jt+1}(b,e)\}_{t\geq 0}$, and joint distributions over assets and productivity levels $\{G_{jt}(b,e)\}_{t\geq 0}$, such that wages are fixed at their steady-state level, firms' optimality conditions and households' optimal consumption-savings decisions are satisfied, and the goods markets clear at the county level:*

$$L_{jt}^{NT} = C_{jt}^{NT} \tag{14}$$

$$L_{jt}^T = \left(\frac{P_j^T}{P^T}\right)^{-\theta} C_t^T \tag{15}$$

Second, we introduce the notion of a *national equilibrium*, where we endogenize national demand for tradables and assets while still requiring that all individual counties remain in the regional equilibrium.

Definition 4 (National equilibrium)**.** *Given initial regional distributions $\{G_{j0}(b,e)\}_{j\in[0,1]}$ and a path for monetary policy $\{r_t\}_{t\geq 0}$, a national equilibrium consists of paths for aggregate quantities $\{\{L_{jt}, L_{jt}^{NT}, L_{jt}^T, C_{jt}, C_{jt}^{NT}, C_{jt}^T, \{C_{jt}^T(j')\}_{j'}\}_{j\in[0,1]}\}_{t\geq 0}$, individual allocation rules $\{\{c_{jt}(b,e), b_{jt+1}(b,e)\}_{j\in[0,1]}\}_{t\geq 0}$, and joint distributions over assets and productivity levels $\{\{G_{jt}(b,e)\}_{j\in[0,1]}\}_{t\geq 0}$, such that wages are fixed at their steady state level, national tradable demand is defined consistently with the paths of regional demands, and the national financial market clears:*

$$C_t^T = \int_0^1 C_{j't}^T dj' \quad \textit{for all } j \tag{16}$$

$$0 = \int_0^1 \left(P_j \int_0^1 b_{jit}\, di \right) dj \tag{17}$$

2.3 Sufficient statistics and sequence-space representation

Before delving into the details of our mechanism, we introduce two objects that will be at the core of our analysis and summarize the dimensions of regional heterogeneity present in the model.

Sequence space representation. We now derive a sequence-space representation of our model (Auclert et al., 2021a). This approach allows us to summarize all regional heterogeneity in households' discount factors, β_j, income risk, e_j, and credit market structure, $\underline{b}_j$, with just two so-called *sequence-space Jacobians*. Throughout the rest of our analysis, we express all sequences in log-deviation from steady state.[8] Because labor unions allocate hours across households, we can express idiosyncratic real income as a function of aggregate county-level quantities only. In particular, we have:

$$z_{ijt} e_{ijt} = \frac{W_{jt} L_{jt}}{P_{jt}} e_{jit}$$

Substituting the expression above into the household's budget constraint (2), it is easy to see that, given the state (b, e), the path of optimal choices $\{c_{jt}(b,e), b_{jt+1}(b,e)\}_{t\geq 0}$ is entirely pinned down by the sequence of aggregate real income $\left\{\frac{W_{jt}}{P_{jt}} L_{jt}\right\}_{t\geq 0} \equiv \{Z_{jt}\}_{t\geq 0}$, together with the sequence of the real interest rate $\{r_t\}_{t\geq 0}$. We can then integrate over states to write aggregate consumption in county j at time t as a function of the sequences of aggregate real income and the real interest rate:

$$\int c_{jt}(b,e) dG_{jt}(b,e) = \mathcal{C}_{jt}\left(\{Z_{js}\}_{s\geq 0}, \{r_s\}_{s\geq 0} \right) \tag{18}$$

Following Auclert et al. (2024b), we denote the partial-equilibrium Jacobian of $\mathcal{C}_{jt}(\cdot)$ with respect to aggregate real labor income $\boldsymbol{Z}_j \equiv (Z_{j0}, Z_{j1}, \dots)'$ by $\boldsymbol{M}_j$, which is a matrix whose element (t,s) is given by $\frac{\partial \ln \mathcal{C}_{jt}(\cdot)}{\partial \ln Z_{js}}$. Similarly, we denote by $\boldsymbol{M}_j^r$ the matrix of partial-equilibrium elasticities of $\mathcal{C}_{jt}(\cdot)$ with respect to the interest rate sequence $\boldsymbol{r} \equiv (r_0, r_1, \dots)'$, that is $\left(\boldsymbol{M}_j^r\right)_{t,s} \equiv \frac{\partial \ln \mathcal{C}_{jt}(\cdot)}{\partial \ln(1+r_s)}$.

[8] In particular, for a generic variable X_{jt}, we denote by $\widehat{\boldsymbol{x}}_j$ the full sequence of log-deviations of variable X_{jt} from its steady state value, i.e., $\widehat{\boldsymbol{x}}_j \equiv \left(\log(X_{j0}) - \log(X_j), \log(X_{j1}) - \log(X_j), \dots\right)'$. For real interest rates r_t, we adopt a slightly different notation and let $\widehat{\boldsymbol{r}} \equiv (\frac{r_0 - r}{1+r}, \frac{r_1 - r}{1+r}, \dots)'$.

Together, $\boldsymbol{M}_j$ and $\boldsymbol{M}^r_j$ capture all household heterogeneity within a given county and thus serve as sufficient statistics for cross-county heterogeneity in iMPCs. The benefit of this approach is that we can be general on what is the driver of regional heterogeneity in iMPCs, whether it is differences in the discount factor, income risk, borrowing limit, or all of the above. Accordingly, we will refer to cross-country heterogeneity in iMPCs as directly capturing heterogeneity in the equilibrium county-level Jacobians $\boldsymbol{M}_j$ and $\boldsymbol{M}^r_j$.

Non-tradable sector share. Similarly, we now define a variable that will capture regional heterogeneity in trade openness: ρ_j, the steady state non-tradable share of labor income.

Definition 5 (ρ_j)**.** *We define ρ_j as county j's non-tradable share of the wage bill in steady state. Formally:*

$$\rho_j = \frac{L_j^{NT} W_j^{NT}}{L_j W_j}$$

Since it represents the share of non-tradable labor income, ρ_j is naturally bounded between 0 and 1. As the following lemma shows, ρ_j captures the extent to which county j is exposed to fluctuations in the non-tradable sector, and thus acts as a sufficient statistic for the transmission of local versus national consumption changes to regional real labor income.

Lemma 1. *In a regional equilibrium, ρ_j governs the pass-through of county j's consumption to employment in county j, and $1 - \rho_j$ governs the pass-through of national tradable consumption to employment in county j:*

$$\widehat{\boldsymbol{\ell}}_j = \rho_j \widehat{\boldsymbol{c}}_j + (1 - \rho_j)\widehat{\boldsymbol{c}}^T \tag{19}$$

where $\widehat{\boldsymbol{c}}^T$ is the log-deviation from steady state of the nationwide demand for tradables, as defined in (16)*.*

Proof. See Appendix A.1. □

Note that this result relies only on the homotheticity of the consumption aggregator, and does not depend on the specific functional forms we assumed. Lemma 1 implies that for the purposes of our analysis we can be very general about what is driving regional heterogeneity in trade openness—whether it is ω_j, capturing demand motives, α_j, governing supply forces, or both. Instead, the key variable we need to track is the steady-state share of non-tradable labor income, ρ_j.[9] Thus, we refer to regional differences in trade openness as meaning regional heterogeneity in ρ_j.

2.4 The regional Keynesian cross

With our measures for county-level household heterogeneity and trade openness established, we can now analyze how a specific county responds to aggregate shocks. A result that has been emphasized by the recent HANK literature is how indirect effects become a major source of the monetary transmission mechanism, once one accounts for household heterogeneity (Kaplan et al., 2018, Bilbiie, 2024, Auclert et al., 2024b). The next proposition shows how these indirect effects—which we refer to as the "regional Keynesian multiplier"—are modified in our regional framework.

Proposition 1 (The regional Keynesian cross)**.** *The first-order impulse response function $\widehat{\boldsymbol{\ell}}_j$ to a monetary shock $\widehat{\boldsymbol{r}}$ and a tradable demand shock $\widehat{\boldsymbol{c}}^T$, which satisfies a regional equilibrium, is given by:*

$$\widehat{\boldsymbol{\ell}}_j = \underbrace{\rho_j \left(\boldsymbol{M}_j^r \widehat{\boldsymbol{r}} + \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}_j \right)}_{\text{Regional equilibrium effects}} + \underbrace{(1-\rho_j)\widehat{\boldsymbol{c}}^T}_{\text{National equilibrium effects}} \tag{20}$$

and the response of regional consumption $\widehat{\boldsymbol{c}}_j$ is given by:

$$\widehat{\boldsymbol{c}}_j = \boldsymbol{M}_j^r \widehat{\boldsymbol{r}} + \rho_j \boldsymbol{M}_j \widehat{\boldsymbol{c}}_j + (1-\rho_j)\boldsymbol{M}_j \widehat{\boldsymbol{c}}^T \tag{21}$$

Proof. See Appendix A.2. □

Proposition 1 characterizes the local employment response in county j in a regional equilibrium, i.e., given a path for the real interest rate $\widehat{\boldsymbol{r}}$, as well as a path for national tradable demand, $\widehat{\boldsymbol{c}}^T$.[10] In particular, (20) shows that the local employment response to a monetary policy shock is fully governed by the two initial impulses, $\widehat{\boldsymbol{r}}$ and $\widehat{\boldsymbol{c}}^T$, and the variables capturing the two dimensions of regional heterogeneity: iMPCs, $\boldsymbol{M}_j$, and intertemporal substitution motives, $\boldsymbol{M}_j^r$, together with the share of the non-tradable labor income, ρ_j.

Before inspecting in greater depth the different channels operating in our regional Keynesian cross, and building intuition for Proposition 1, it is useful to briefly consider two limit cases.

[9] Under sticky wages, an additional variable is needed for trade openness, namely the steady-state share of non-tradable consumption. See Appendix B for a detailed discussion.

[10] In the next section, the path for $\widehat{\boldsymbol{c}}^T$ will be further endogenized in the national equilibrium

Corollary 1. *When $\rho_j \to 1$, the first-order response of employment $\widehat{\boldsymbol{\ell}}_j$ to a monetary shock $\widehat{\boldsymbol{r}}$, which satisfies a regional equilibrium, solves a standard intertemporal Keynesian cross (Auclert et al., 2024b):*

$$\widehat{\boldsymbol{\ell}}_j \quad = \quad \boldsymbol{M}_j^r \widehat{\boldsymbol{r}} + \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}_j \quad = \quad \widehat{\boldsymbol{c}}_j \tag{22}$$

Corollary 1 considers the limit case in which county j is a fully closed economy within a monetary union. In this scenario, intuitively, the local employment response to a monetary shock is not going to be affected by fluctuations in the tradable sector. Hence, it follows the standard closed-economy intertemporal Keynesian cross as described in Auclert et al. (2024b), which our framework nests as a limit case.

Corollary 2. *When $\rho_j \to 0$, the first-order response of employment $\widehat{\boldsymbol{\ell}}_j$ to a monetary shock $\widehat{\boldsymbol{r}}$ satisfying a regional equilibrium does not depend on j's characteristics, and is fully nationally determined :*

$$\widehat{\boldsymbol{\ell}}_j \quad = \quad \widehat{\boldsymbol{c}}^T \tag{23}$$

and the response of regional consumption $\widehat{\boldsymbol{c}}_j$ is given by:

$$\widehat{\boldsymbol{c}}_j \quad = \quad \boldsymbol{M}_j^r \widehat{\boldsymbol{r}} + \boldsymbol{M}_j \widehat{\boldsymbol{c}}^T \tag{24}$$

Corollary 2 shows how, in the limit case in which a county faces no trade frictions and is thus fully exposed to national business cycles, the county-level response of *employment* to monetary policy shocks is fully independent of county-specific characteristics. In particular, county-specific iMPCs, as summarized by $\boldsymbol{M}_j$, do not matter in determining the county-specific employment response.[11] This result is going to be useful later on, once we derive the *national* Keynesian cross.

Decomposing the channels. Proposition 1 provides a decomposition of the local employment response to a monetary policy shock into two channels, each operating either via the tradable goods market or via the non-tradable goods market.

Regional Equilibrium Effects—The first term in (20) captures the effect of county j's exposure to local fluctuations on the regional employment response. This regional exposure channel is premultiplied by ρ_j because this factor exactly captures the pass-through of fluctuations in the non-tradable sector, i.e., local fluctuations, to the local economy. Because

[11] The path of the consumption response is still going to be shaped by households' iMPCs.

$\rho_j \in [0, 1]$ the presence of regional trade flows dampens the standard closed-economy direct and indirect channels of monetary policy. First, a monetary policy expansion induces a rise in consumption according to the matrix of intertemporal elasticity of substitution $\boldsymbol{M}_j^r$—this is the standard interest rate channel of monetary policy. This rise in consumption only affects local employment in the non-tradable sector, and hence employment rises by $\rho_j \boldsymbol{M}_j \widehat{\boldsymbol{r}}$. Second, indirect effects are also dampened by the presence of tradable goods. The rise in households' demand, induced by the interest rate change, fuels an increase in households' real income, which in turn translates into additional consumption depending on the matrix of intertemporal marginal propensities to consume, $\boldsymbol{M}_j$. Hence, this further growth in consumption generates an increase in employment. However, from a regional perspective, this propagation mechanism only takes place in the non-tradable sector. For this reason, in our framework the regional Keynesian multiplier channel is captured by the term $\rho_j \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}_j$.

National Equilibrium Effects—Besides dampening the standard closed-economy channels of monetary transmission, the presence of tradable goods also introduces new propagation mechanisms. These are represented by the second term in (20), which captures the role played by county j's exposure to national fluctuations on the local employment response. This national exposure channel is premultiplied by a factor $1 - \rho_j$ because, as discussed before, national fluctuations in the demand for the tradable good are passed-through to the local economy only via the tradable sector. This national exposure channel enters the regional Keynesian cross as a shifter, thus feeding the regional Keynesian multiplier but not augmenting it. This comes from the fact that regions are atomistic and, thus, their local response does not matter for nationwide tradable demand. In the next section, we will show how the national tradable response becomes an endogenous object, giving rise to the national Keynesian cross.

The next proposition derives the general solution to (20) in a regional equilibrium:

Proposition 2. *The first-order response of employment $\widehat{\boldsymbol{\ell}}_j$ to a monetary shock $\widehat{\boldsymbol{r}}$ and a tradable demand shock $\widehat{\boldsymbol{c}}^T$ is given by:*

$$\widehat{\boldsymbol{\ell}}_j = \underbrace{\left(\boldsymbol{I} - \rho_j \boldsymbol{M}_j\right)^{-1}}_{\text{Keynesian multiplier}} \left(\underbrace{\rho_j \boldsymbol{M}_j^r \widehat{\boldsymbol{r}}}_{\text{Interest rate channel}} + \underbrace{(1 - \rho_j) \widehat{\boldsymbol{c}}^T}_{\text{Tradable demand}} \right) \tag{25}$$

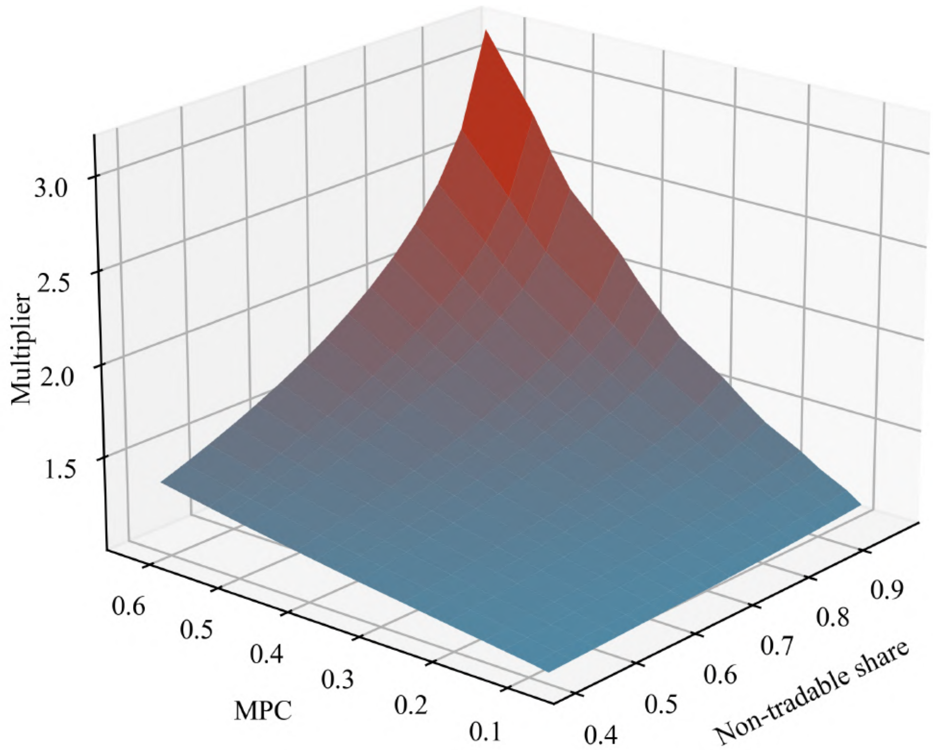
Proposition 2 showcases how at the local level the transmission of monetary policy is jointly affected by our two dimensions of regional heterogeneity. This is particularly important when it comes to the regional multiplier matrix $\boldsymbol{\mathcal{M}}_j \equiv \left(\boldsymbol{I} - \rho_j \boldsymbol{M}_j\right)^{-1}$. First, by its very nature, the multiplier term is non-linear both in the demand-side channels at work in the standard HANK framework—captured by the iMPC matrix $\boldsymbol{M}_j$—and the supply-side channels of the standard open-economy framework, which are captured by the non-tradable share, ρ_j. Second, and more importantly, these two channels interact in shaping the magnitude of the multiplier, acting as complements. This means that the marginal effect of iMPCs on the regional Keynesian multiplier is increasing in ρ_j, and vice-versa. The intuition for this is that regions with low ρ_j have a small exposure to local fluctuations to begin with. Hence, the role that iMPCs play in propagating shocks via Keynesian multiplier effects is less important for these regions. Similarly, the Keynesian multiplier channel is small in counties with low iMPCs, thus making the extent to which these counties are exposed to it less relevant.

Figure 1 illustrates this mechanism. It plots the first entry of the multiplier matrix $\boldsymbol{\mathcal{M}}_j$ across different values of MPCs and ρ_j.[12] One important takeaway from Figure 1 is that the non-linearities between ρ_j, MPC_j, and the multiplier start kicking in for relatively large values of both ρ_j and MPC_j. Quantitatively, this means that it is important that our model MPCs are in line with the relatively large ones measured in the data. This is one of the reasons why we favour our HANK approach over a standard RANK framework. Moreover, as we show in Section 3, our empirical measure of ρ_j approaches unity in a fairly sizeable share of counties. As a result, the non-linearity of the multiplier will play a key role in shaping the national response to monetary policy, as the joint spatial distribution of MPCs and openness to trade will govern the national pass-through of monetary policy. This will be the subject of the next section.

Finally, note that this iMPC-trade openness complementarity is not limited to the regional multiplier term. Similarly to the case of the Keynesian multiplier, the magnitude of the local direct effect of monetary policy—the term $\rho_j \boldsymbol{M}_j^r$ in (25)—is also shaped by the interaction between openness to trade and intertemporal substitution motives. Hence, in our framework, local openness to trade governs the extent to which iMPCs and intertemporal substitution motives—and, more generally, local household heterogeneity—matter for the local employment response to monetary policy. In Section 3 we present empirical support for this complementarity in the context of the US economy.

[12]In the graph, we equate our MPC measure to the first entry of the $\boldsymbol{M}_j$ matrix, which captures the on-impact response to an unanticipated increase in real income. This is consistent with our approach in the following sections of the paper, where we match the high-dimensional object $\boldsymbol{M}_j$ to the data using its first entry $\boldsymbol{M}_j[0,0]$ as a measure of the static MPC.

Figure 1: MPC-trade openness complementarity and the regional Keynesian multiplier



Note: the vertical axis represents the first entry of the multiplier matrix $\boldsymbol{\mathcal{M}}_j[0,0]$, the MPC axis is the first entry of the iMPC matrix $\boldsymbol{M}_j[0,0]$, and the non-tradable share axis represents ρ_j. Multipliers are computed on the basis of the baseline calibration (described in Section 4.2), taking national prices as exogenous, varying the discount rate β_j to span different levels of MPCs, and varying α_j to span different ρ_j's. County-level prices respond endogenously so that county-level markets clear.

2.5 The national Keynesian cross

Our regional Keynesian cross result characterizes the response to a monetary shock in the cross-section of *regions* and is therefore useful to understand how our two dimensions of regional heterogeneity shape the heterogeneous response to monetary shocks across counties. We now study the impact of regional heterogeneity on the *nationwide* propagation of monetary policy. To do so, we aggregate our continuum of regional Keynesian crosses (20) and endogenize the path of national demand for tradables by using our national equilibrium concept.

Proposition 3 (The national Keynesian cross)**.** *The first-order nationwide response of consumption $\widehat{\boldsymbol{c}}$ to a monetary shock $\widehat{\boldsymbol{r}}$ satisfying a national equilibrium is characterized by:*

$$\widehat{\boldsymbol{c}} = \underbrace{\Big(\boldsymbol{M} + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, (1-\rho_j)\boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j)\Big)\widehat{\boldsymbol{c}}}_{\textit{National multiplier}} + \underbrace{\Big(\boldsymbol{M}^r + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j^r)\Big)\widehat{\boldsymbol{r}}}_{\textit{National interest rate channel}} \tag{26}$$

and the national employment response $\widehat{\boldsymbol{\ell}}$ is given by:

$$\widehat{\boldsymbol{\ell}} = \Big(\boldsymbol{M} + \mathrm{Cov}(\rho_j, \boldsymbol{M}_j)\Big)\widehat{\boldsymbol{\ell}} \, + \, \Big(\boldsymbol{M}^r + \mathrm{Cov}(\rho_j, \boldsymbol{M}_j^r)\Big)\widehat{\boldsymbol{r}} \, + \, \mathrm{Cov}((1+\rho_j-\rho))\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j) \tag{27}$$

Where $\boldsymbol{\mathcal{M}}_j \equiv \left(\boldsymbol{I} - \rho_j \boldsymbol{M}_j\right)^{-1}$, $\widehat{\boldsymbol{\ell}}_j$ is given by (20), $\widehat{\boldsymbol{\ell}} \equiv \mathbb{E}\,\widehat{\boldsymbol{\ell}}_j$, $\widehat{\boldsymbol{c}} \equiv \mathbb{E}\,\widehat{\boldsymbol{c}}_j$, $\rho \equiv \mathbb{E}\,\rho_j$, $\boldsymbol{M} \equiv \mathbb{E}\,\boldsymbol{M}_j$, and $\boldsymbol{M}^r \equiv \mathbb{E}\,\boldsymbol{M}_j^r$, where $\mathbb{E}$ and Cov *are, respectively, the expectation and covariance operators weighted by steady state tradable consumption.*

Proof. See Appendix A.3. □

Proposition 3 highlights the impact of regional heterogeneity on the aggregate transmission of monetary policy. The result in (26) delivers a closed-form solution for the nationwide consumption response to the monetary impulse. Similarly, the result obtained in (27)—together with (20)—provides a fixed point characterizing the aggregate employment response. As already discussed in the context of Figure 1, the non-linearity present in the regional Keynesian multiplier implies that regional heterogeneity does not wash out in the aggregate and, thus, regional distributions matter for the national response. In particular, Proposition 3 shows how the key object summarizing the role of regional heterogeneity for the aggregate response to shocks is the cross-sectional covariance between (a convolution of) iMPCs and the share of non-tradable labor income. The reason why it is covariances, rather than correlations, goes again back to Figure 1. Because of the non-linear nature of the regional Keynesian multiplier, mean-preserving spreads in either MPCs or the shares of non-tradable income are going to change the average response. Hence, the variance of these two objects matters.

Besides that, the fact that the interaction between $\boldsymbol{M}_j$ and ρ_j shapes the magnitude of the multiplier, as explained before, implies that the correlation between the two matters as well. Together, this implies that we need to keep track of covariances, i.e., the sorting between iMPCs and non-tradable employment across regions. Proposition 3 also makes clear that the way the joint distribution of non-tradable employment and iMPCs across space affects the national response is by modifying the national sequence-space Jacobians $\boldsymbol{M}$ and $\boldsymbol{M}^r$. Accordingly, a key implication of this is that two economies which are identical on average, i.e., share the same average Jacobians $\boldsymbol{M}$ and $\boldsymbol{M}^r$, can display very different responses to shocks if they feature different regional distributions around this average.

We can also leverage Proposition 3 to analyze the role of regional heterogeneity for the nationwide transmission of monetary policy under interesting limiting cases. Importantly, this allows us to derive a neutrality result under which regional heterogeneity in iMPCs is irrelevant for the aggregate transmission of monetary shocks in a monetary union.

Corollary 3 (Limit cases and a regional "as-if" benchmark)**.** *When ρ_j is constant across space and equal to ρ we have:*

$$\widehat{\boldsymbol{c}} = \Big(\boldsymbol{M} + \rho(1-\rho)\,\mathrm{Cov}(\boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j)\Big)\widehat{\boldsymbol{c}} + \Big(\boldsymbol{M}^r + \rho\,\mathrm{Cov}(\boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j^r)\Big)\widehat{\boldsymbol{r}}$$

and similarly for employment:

$$\widehat{\boldsymbol{\ell}} = \boldsymbol{M}\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + \mathrm{Cov}(\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j)$$

In the limit case of $\rho_j \to 0$ for all j, heterogeneity in iMPCs across regions does not matter for the nationwide employment response:

$$\widehat{\boldsymbol{c}} = \widehat{\boldsymbol{\ell}} = \boldsymbol{M}\,\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\,\widehat{\boldsymbol{r}} \tag{28}$$

Corollary 3 follows directly from Proposition 3. Equation (28) is a neutrality result for regional heterogeneity, in the spirit of Werning (2015). It states that in the limit case in which all goods are fully tradable, only average nationwide iMPCs matter for the national response, while *heterogeneity* in iMPCs across regions is irrelevant.

Moreover, recall that Corollary 2 already showed how when $\rho_j \to 0$ the employment response to a monetary shock is equalized across regions. Thus, Corollaries 2 and 3 jointly imply that in the limiting case where all goods are tradable, the only factors that matter for both the national and regional employment responses are the nationwide *average* iMPC Jacobians $\boldsymbol{M}$ and $\boldsymbol{M}^r$, while all regional heterogeneity around this average is irrelevant. In this scenario, the response of a monetary union comprising multiple regions can be accurately approximated by that of a standard single-region economy. However, whenever trade frictions between regions are present, a representative-region model that ignores heterogeneity in iMPCs will lead to misspecified predictions for the national response to monetary shocks.

In summary, this section has demonstrated how heterogeneity in regional iMPCs and non-tradable employment shapes monetary policy transmission within a multi-region HANK framework. We have characterized the full general-equilibrium economic responses to monetary policy shocks both at the level of an individual county and at the level of the monetary union. In the latter case, the joint distribution of iMPCs and non-tradable shares determines the aggregate macroeconomic response. Additionally, we have derived sufficient conditions for an "as-if" benchmark, in which regional heterogeneity is irrelevant for aggregate dynamics, allowing the economy to be approximated by a representative region. In the following sections, we argue both empirically and quantitatively that deviations from this benchmark are considerable.

3 Empirical analysis

In this section, we turn to U.S. micro-data and construct empirical counterparts to our model's key variables: county-level marginal propensities to consume (MPCs) and non-tradable employment shares. We then document substantial heterogeneity in the county-level employment response to monetary policy shocks. Finally, we demonstrate that the geography of local MPCs and non-tradable employment systematically explains this heterogeneity.

3.1 Data sources

Employment. Our main data source is the Local Area Unemployment Statistics (LAUS) database from the Bureau of Labor Statistics. The LAUS register is a non-survey-based dataset which combines multiple data sources to provide monthly employment estimates for different levels of regional disaggregation. In what follows, we focus on county-level employment.[13]

We obtain annual county-sector-level employment information for 4-digit North American Industry Classification System (NAICS) industries from the County Business Patterns (CBP) dataset published by the U.S. Census. Prior to 1998, data is based on the SIC industry classification. Hence, we link SIC sectors to NAICS according to the SIC-NAICS concordance tables provided by the U.S. Census. We then classify 4 digit-NAICS sectors into tradable and non-tradable industries according to the retail and world trade based definition of Mian and Sufi (2014).

Monetary policy. In order to capture monetary policy surprises, we follow the high-frequency identification approach.[14] Specifically, as our baseline instrument for monetary shocks we rely on the Jarociński and Karadi (2020) shock series. These shocks are based on the change in the 3-month ahead Fed Funds futures within a 30-minute window around FOMC announcements, adjusted for the Fed's information effect.[15] For robustness, we also consider the standard, unadjusted, Gertler and Karadi (2015) high-frequency shock,

[13] As of 2020, there were 3,143 counties across the 50 U.S. states. Our dataset comprises a total of 3,140 counties, 92.50% of which are present in all months of the sample.

[14] See, among others, Kuttner (2001), Gürkaynak et al. (2005), Gertler and Karadi (2015), Gorodnichenko and Weber (2016), Nakamura and Steinsson (2018).

[15] Specifically, we use the poor man's shock series.

as well as the narrative instrument proposed in Romer and Romer (2000) and updated by Miranda-Agrippino and Rey (2020). Throughout the rest of our analysis, we normalize the sign of the measure of monetary shocks ε_t such that positive values are associated with expansionary shocks. Moreover, we also normalize ε_t to have a standard deviation of unity.

Data for regional MPCs. To construct our regional measure of MPCs we rely on two data sources. The first consists of four special survey modules designed by Fuster et al. (2020) and fielded as part of the NY Fed Survey of Consumer Expectations (SCE). These modules are constructed to elicit respondents' self-reported MPCs out of windfall gains and losses of different magnitudes. A total of 2,586 panelists participated in the survey across four waves between March 2016 and March 2017. We specifically rely on a question asking respondents to report their quarterly MPC out of a windfall loss of 500$.

Second, we use the American Community Survey (ACS) 5-year data to obtain county-level information on the socio-demographic characteristics of the local population. Specifically, we extract data on the number of households in each county, categorized by the age, race, educational attainment, and employment status of the householder, as well as by household income, homeownership status, and whether the household has non-zero capital income. This data is available at an annual frequency and covers the period from 2011 to 2021.

To validate our novel regional MPC measure, we also use existing county-level data on stock market wealth from Chodorow-Reich et al. (2021). This measure of wealth is obtained by applying an improved version of the canonical capitalization method to data on taxable dividend income aggregated at the county level. We construct an annual measure of stock market wealth per capita, spanning the years 1989-2015, by using county population data from the U.S. Census.

3.2 The geography of MPCs and non-tradable employment

Constructing regional MPCs. Marginal propensities to consume are a challenging object to estimate in the data. First, because identified exogenous variations in households' income are rare. Second, because it is often difficult to get access to household-level expenditure data at a sufficiently high frequency. Moreover, our research question requires an accurate proxy for MPCs that varies granularly across space. However, since most estimates of MPCs rely on survey data, the sample size is usually not large enough to construct accurate estimates at a fine local level.

We address these challenges by extending the method proposed in Patterson (2023) to a regional setting. In doing so, we construct a novel measure of county-level MPCs for the United States using a two-step approach. In the first step, drawing on data from Fuster et al. (2020), we run a respondent-level regression of self-reported MPCs on indicator variables representing respondents' age, race, household income, educational attainment, employment status, homeownership status, and an indicator variable that equals one if the respondent reports strictly positive financial income and zero otherwise.[16] Our estimation specification reads as follows:

$$\begin{aligned} MPC_i = \alpha &+ \sum_{s=1}^{4} \beta_s^A D_{si}^A + \sum_{s=1}^{5} \beta_s^R D_{si}^R + \sum_{s=1}^{9} \beta_s^Y D_{si}^Y + \sum_{s=1}^{5} \beta_s^E D_{si}^E + \sum_{s=1}^{3} \beta_s^U D_{si}^U + \sum_{s=1}^{2} \beta_s^H D_{si}^H \\ &+ \sum_{s=1}^{2} \beta_s^K D_{si}^K + u_i \end{aligned} \tag{29}$$

where α is a constant and the variables D_{si}^A to D_{si}^K are indicator variables for race, age, income, educational attainment, employment status, homeownership and financial income, respectively.[17]

Next, from the ACS data we compute county-level shares of households belonging to each of the groups included in (29). Because of the additive structure of (29) we can then simply compute county-level MPCs as a weighted average of our estimates from (29):

$$\begin{aligned} \widehat{MPC}_{jt} = &\hat{\alpha} + \sum_{s=1}^{4} \hat{\beta}_s^A \mu_{sjt}^A + \sum_{s=1}^{5} \hat{\beta}_s^R \mu_{sjt}^R + \sum_{s=1}^{9} \hat{\beta}_s^Y \mu_{sjt}^Y + \sum_{s=1}^{5} \hat{\beta}_s^E \mu_{sjt}^E + \sum_{s=1}^{3} \hat{\beta}_s^U \mu_{sjt}^U + \sum_{s=1}^{2} \hat{\beta}_s^H \mu_{sjt}^H \\ &+ \sum_{s=1}^{2} \hat{\beta}_s^K \mu_{sjt}^K \end{aligned} \tag{30}$$

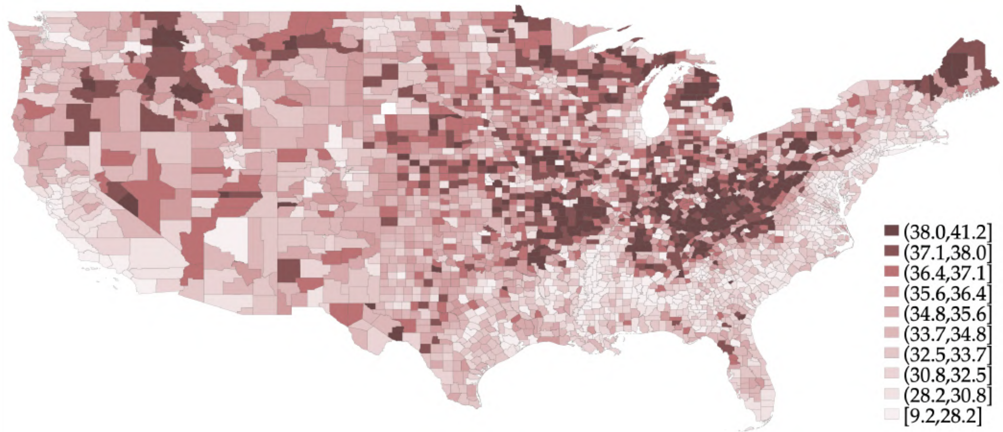
Where $\mu_{sjt}^g \equiv \frac{N_{sjt}^g}{N_{jt}}$ is the fraction of households in county j and year t that belong to class s for the socio-economic characteristic g.

Importantly, because (29) relies on bins and is hence non-parametric in nature, our methodology does not impose any linearity and does not rely on average measures when estimating MPCs. On the contrary, we account for the full distribution of households

[16] In particular, we include 5 race categories (white, black, Asian, hispanic, and other), 4 age categories (less than 25, 25-44, 45-64, and more than 65 years old), 10 categories for household income (less than 10k, 10-20k, 20-30k, 30-40k, 40-60k, 60-75k, 75-100k, 100-150k, 150-200k, and more than 200k USD), 5 categories for educational attainment (less than high school, high school diploma, some college or finished college, bachelor degree, masters degree or more), and 3 categories for employment status (working full- or part-time, not working but would like to work, and other).

[17] The estimates from (29) are reported in Table C.3 in Appendix C.

Figure 2: The geography of MPCs in the United States



Note: This figure plots the geographic distribution of of U.S. county-level MPCs. MPCs are averaged over the years 2011-2019. See main text for details on the measurement procedure.

along both economic and socio-demographic dimensions. This is consistent with the theoretical insights of heterogeneous-agent models, which highlight that the cross-sectional average MPC is not well approximated by the MPC of the average household, since the relationship between an individual's MPC and her states is, in general, non-linear. Our procedure is also flexible and portable to different contexts.

Figure 2 shows the map of estimated quarterly regional MPCs, averaged over the years 2011-2019. Our estimates range from 0.09 to 0.41, with a population-weighted average of 0.30 and a standard deviation of 0.04.[18] Counties in rich metropolitan areas such as the New York, Chicago, Miami, and San Francisco area have low MPCs, while counties in the Appalachian coal and mining region and the rust-belt area tend to display large MPCs.

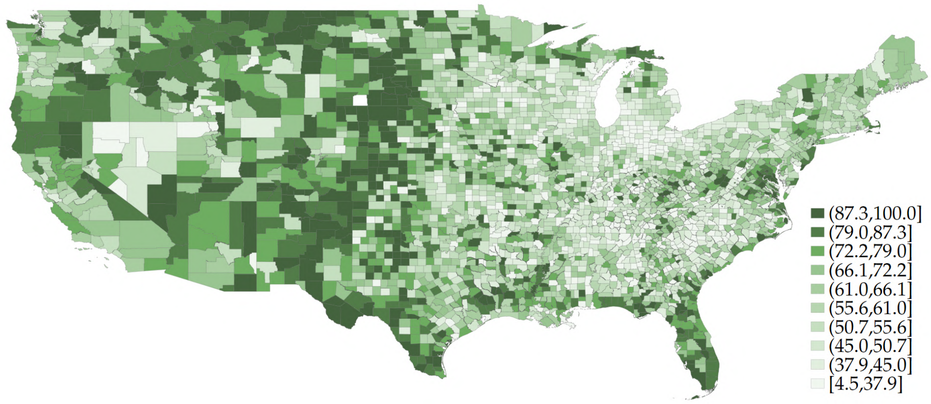
Relative to the existing empirical literature, we observe that our estimate of the average MPC is in the ballpark of standard estimates but on the high end of the spectrum.[19] In particular, across various studies, the reported average quarterly MPC out of transitory income changes is between 15% and 30% (Jappelli and Pistaferri, 2010, Havranek and Sokolova, 2020). However, the average masks substantial cross-sectional heterogeneity in MPCs (Misra and Surico, 2014, Fagereng et al., 2021, Lewis et al., 2024, Boehm et al., 2025). Our contribution is to measure and quantify precisely the *spatial* dimension of this heterogeneity, which is an underexplored aspect of this recent empirical literature.

As a validation check of our methodology and results, we correlate our regional measure of MPCs with the (logarithm of) stock market wealth per capita from Chodorow-

[18]See Figure C.8 in Appendix C for a histogram of the distribution of regional MPCs in 2016.

[19]This comes from the fact that the average self-reported MPC among the respondents to the Fuster et al. (2020) survey question eliciting MPC out of 500$ windfall losses is 0.3.

Figure 3: The geography of non-tradable employment in the United States



Note: This figure plots the geographic distribution of U.S. county-level non-tradable employment shares which are averaged over the years 2011-2019. See the main text for details on the empirical procedure and definitions.

Reich et al. (2021). A good measure of MPCs should be negatively associated with wealth. Importantly, note that we have never relied upon the stock market wealth data in our MPC computation approach. Thus, we are comparing two measures that are built from very different underlying microeconomic data. As shown on Figure C.9 in Appendix C, the two measures have a population-weighted correlation coefficient of -0.49, which is statistically significant at the 1% level. Intuitively, to the extent that stock market wealth captures total wealth, it is reassuring and theoretically consistent that the relationship between wealth and our MPC measure is strongly negative.

County-level non-tradable employment share. We now compute the empirical counterpart to our theory-based measure of the share of non-tradable income ρ_j. To do so, we leverage data from the CBP and classify tradable and non-tradable 4-digit industries using the Mian and Sufi (2014) retail and world-trade based classification. For each county j and year t, we define our baseline trade openness variable as the non-tradable to tradable employment ratio $\tilde{\rho}_{jt} \equiv L_{jt}^{NT}/(L_{jt}^{NT} + L_{jt}^{T})$, where L_{jt}^{NT} represents the total number of people working in non-tradable sectors in county j and year t, while L_{jt}^{T} is the total number of people employed in tradable sectors in the same county-year unit.[20]

Figure 3 presents the resulting spatial distribution of $\tilde{\rho}_{jt}$ across counties, averaged over all years in our sample. The values of $\tilde{\rho}_{jt}$ range from 0.045 to 1, with an average of 0.6 and a standard deviation of 0.16. What is the relationship between our two key spatial

[20]We define our $\tilde{\rho}$ measure based on employment data rather than payroll data due to better coverage at the 4-digit NAICS industry classification level. Furthermore, since the industry classification from Mian and Sufi (2014) does not encompass all industries, we define our denominator as the sum of tradable and non-tradable employment to ensure closer alignment between our empirical and theoretical definitions.

characteristics? Figure C.10 in Appendix C shows that regional MPCs and non-tradable employment shares are mildly negatively correlated, with a population-weighted correlation coefficient of -0.31. This occurs because rich metropolitan areas, everything else equal, tend to feature larger shares of non-tradable employment.

Taking stock, we have so far constructed the empirical counterparts to our two theory-driven measures of regional MPCs and non-tradable employment shares. We have also documented large regional heterogeneity in both local MPCs and non-tradable employment shares across US counties. We now turn to analyze the implications of this regional heterogeneity and in particular its ability to explain the geography of local, county-level responses to monetary policy shocks.

3.3 Regional effects of U.S. monetary policy shocks

We now estimate county-level employment responses to US monetary shocks. To do so, we estimate panel lag-augmented local projections (Jordà, 2005, Montiel Olea and Plagborg-Møller, 2021).[21] In particular, for each county j and month t and for horizons $h = 0, \ldots, 36$, we run the following regression:

$$\Delta \ln(L_{j,t+h}) = \alpha_{jh} + \delta_{th} + \beta_{jh} \times \varepsilon_t + \sum_{\ell=1}^{12} \gamma_{h\ell} \Delta \ln(L_{j,t-\ell}) + u_{jht} \tag{31}$$

where $\Delta \ln(L_{j,t+h}) = \ln(L_{j,t+h}) - \ln(L_{j,t-1})$ represents the h-month ahead cumulative change in employment in county j, α_{jh} is a county fixed effect, and δ_{th} is a time fixed effect. Finally, $\Delta \ln(L_{jt-\ell}) = \ln(L_{j,t-1}) - \ln(L_{j,t-\ell-1})$ denotes past county-level employment growth, while ε_t is the monetary surprise.

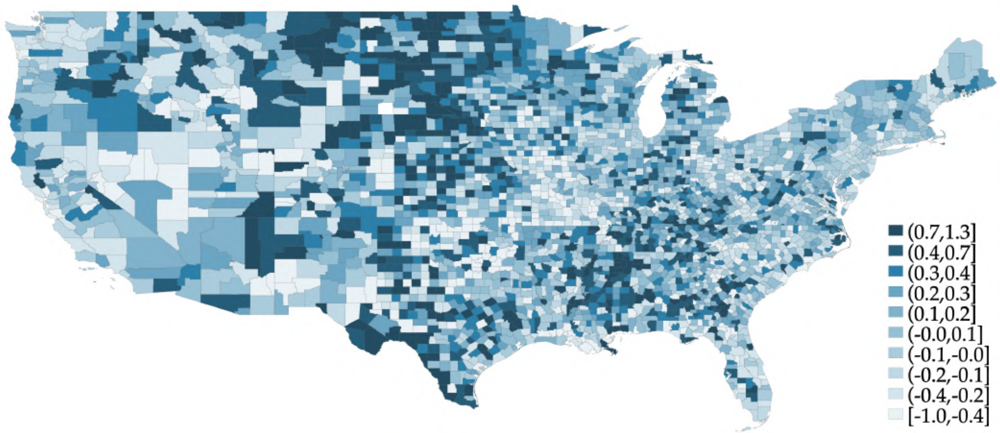
Figure 4 presents a map of the county-specific estimated coefficients $\hat{\beta}_{j,h=36}$ for a 3-year ahead horizon, $h = 36$. These estimates are centered around zero, as they represent deviations from the population-weighted nationwide average response.[22] The figure illustrates substantial cross-county heterogeneity in local employment responses to the monetary shock. Specifically, some counties experience employment increases up to 1.3 percentage points higher than the average response, while others show a lower response of up to 1 percentage point relative to the average county.[23] Notably, this heterogeneity does not appear to be randomly distributed across regions. On the contrary, the responses exhibit

[21] All regression results are weighted by county population as of the year 2010.

[22] Figure C.11 in Appendix C plots the associated histogram. The choice of horizon is immaterial: from 12 months to 36 months the estimated responses are highly correlated, as is shown in Figure C.12 in Appendix C.

[23] For reference, the estimate for the 36-month ahead average response is 0.2 percent.

Figure 4: Regional heterogeneity in the effects of U.S. monetary policy



Note: This figure plots the 3-year ahead county-specific cumulative employment responses to a 1 standard deviation expansionary monetary policy shock, $\hat{\beta}_{j,36}$, estimated from the panel local projection (31). Estimates are in terms of percentage-point deviations from the (population-weighted) average response.

systematic geographic clustering and may therefore be explained by local characteristics. Below, we show that local MPCs and openness to trade turn out to be two important drivers of regional differences in the response to monetary shocks.

Monetary policy, MPCs, and openness to trade. We now investigate how our two theory-based measures of local MPCs and openness to trade shape the regional heterogeneity in the response to monetary policy. To do so, we rank counties in quintiles according to our two variables of interest: MPC_j, and the non-tradable employment ratio, $\tilde{\rho}_j$.[24] We then construct two indicator variables: D_j^{ρ}, which equals one when the ratio of non-tradable to tradable employment $\tilde{\rho}_j$ in county j is in the bottom 80% of the cross-section of counties; and D_j^{M}, which equals one for counties in the bottom 80% of the MPC distribution. We then run the following lag-augmented panel local projection (Montiel Olea and Plagborg-Møller, 2021), with standard errors two-way clustered at the time and county level (Almuzara and Sancibrián, 2024):

$$\Delta \ln(L_{jt+h}) = \underbrace{\alpha_{jh} + \delta_{th}}_{\text{Fixed effects}} + \underbrace{\beta_h^{\rho} \times D_j^{\rho} \times \varepsilon_t}_{\text{Openness interaction}} + \underbrace{\beta_h^{M} \times D_j^{M} \times \varepsilon_t}_{\text{MPC interaction}} + \underbrace{\beta_h^{M\times\rho} \times D_j^{\rho} \times D_j^{M} \times \varepsilon_t}_{\text{Triple interaction}}$$
$$+ \underbrace{\sum_{\ell=1}^{12} \gamma_{h\ell} \Delta \ln(L_{jt-\ell})}_{\text{Lagged controls}} + u_{jht}$$

[24] Our ρ_j measure is average across our full 1990-2019 sample. Our regional estimates for MPCs are only available starting from 2011, so the average is based on the time window 2011-2019.

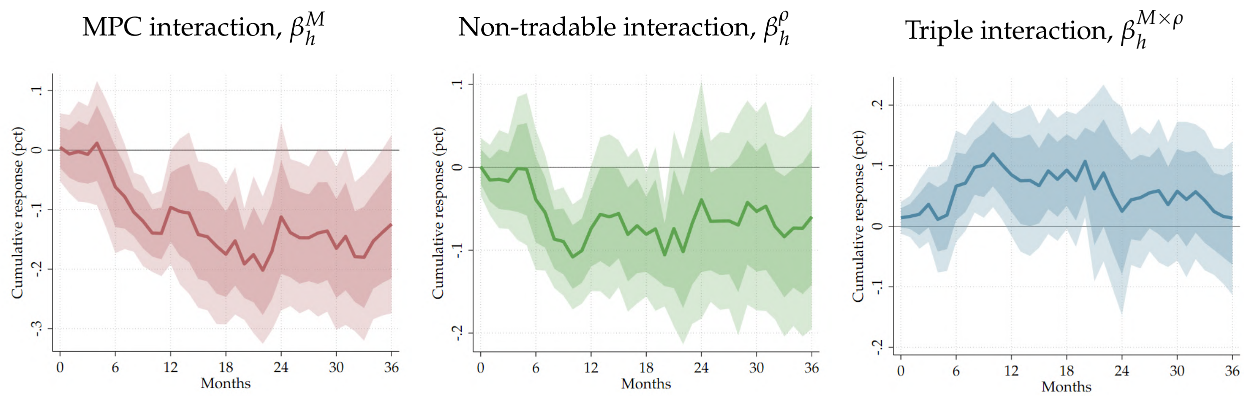
where the definitions of $\Delta \ln(L_{jt+h})$, ε_t, and $\Delta \ln(L_{jt-\ell})$ are the same as in (31), α_{jh} is a county fixed-effect, δ_{th} is a time fixed effect, and D_j^{ρ} and D_j^{M} are the indicator variables defined above.

Note that while the time fixed effect δ_{th} absorbs the monetary shock, our focus is on the *differential* response to the shock across counties. Therefore, (32) crucially includes interaction terms between the monetary shock and our newly constructed indicator variables. Because we are interacting the shock with binary variables, the interpretation of the coefficients is straightforward: the baseline group is represented by counties which are in the top quintile of the non-tradable to tradable employment ($D_j^{\rho} = 0$) and in the top quintile of the MPC ($D_j^{M} = 0$) distributions. Then, β_h^{ρ} simply represents the differential response of low non-tradable employment, high MPC counties (for which $D_j^{\rho} = 1$ and $D_j^{M} = 0$) relative to the baseline group. Similarly, β_h^{M} represents the differential response of low MPC, high non-tradable employment counties (for which $D_j^{M} = 1$ and $D_j^{\rho} = 0$) relative to the baseline. Finally, the response of low MPC, low non-tradable employment counties (for which $D_j^{M} = 1$ and $D_j^{\rho} = 1$) relative to the baseline group is given by the sum of the three coefficients $\beta_h^{\rho} + \beta_h^{M} + \beta_h^{M\times\rho}$. Thus, the coefficient $\beta_h^{M\times\rho}$ captures the interaction effect of simultaneously having both low MPC and low non-tradable employment, over and above the combined individual effects of each variable.

Figure 5 plots impulse responses with respect to a 1 standard deviation expansionary monetary shock, ε_t. The central panel shows the estimates of $\hat{\beta}_h^{\rho}$, along with 68% and 90% confidence bands. Compared to counties for which $D_j^{\rho} = 0$ and $D_j^{M} = 0$, counties which are in the bottom 80% of the non-tradable to tradable employment distribution tend to respond less to monetary shocks. In particular, following an expansionary shock, these regions experience a cumulative increase in employment around 0.1% smaller relative to the baseline group. In other words, we document that high-$\tilde{\rho}_j$ regions are more reactive to monetary impulses. To put this estimate in perspective, consider that in our sample a 1 standard deviation expansionary monetary shock corresponds to roughly a 10 basis points cut in the Fed funds rate and leads to a 3-year ahead cumulative employment response of 0.29%. Thus, besides being statistically significant, our estimated *differential* response of 0.1% appears economically sizeable, when compared to the average effect.

Similarly, the left panel of Figure 5 displays the estimated $\hat{\beta}_h^{M}$. We find that low-MPC counties experience a smaller employment response to monetary shocks, compared to high-MPC areas. In particular, the cumulative change in employment is, on average, roughly 0.2% smaller for those counties for which $D_j^{M} = 1$, relative to the baseline. Thus, we find that high-MPC regions are more reactive to monetary impulses. Again, this magnitude appears economically large when compared to the average employment effect of

Figure 5: Heterogeneous responses by local MPC and non-tradable employment share



Note: impulse response functions to a 1 standard deviation expansionary monetary shock. Lightly (darkly) shaded areas represent 90% (68%) confidence bands. Standard errors are two-way clustered at the time and county levels. The y-axes represent the cumulative percentage change in employment. The x-axes represent months elapsed since the shock.

the same monetary shock.

Finally, we report estimates for the interaction term, $\hat{\beta}_h^{M\times\rho}$, which captures complementarities between MPCs and the non-tradable employment share. These are shown on right panel of Figure 5. We find that the interaction term is positive, statistically and economically significant. This means that when both the MPC and non-tradable employment share variables are low (when the respective dummies take the value of 1), the average elasticity of local employment to a monetary shock is *higher* than what would be predicted by the sum of the marginal effects of low MPC and low non-tradable share. In other words, the marginal effect of MPC (non-tradable share) on the response to monetary shocks is increasing in the non-tradable share (MPC). Recall that our theory—as has been discussed extensively throughout Section 2.4 and also visualized on Figure 1—predicts precisely this positive complementary between MPCs and trade openness. In our model, the regional Keynesian multiplier is the largest for regions characterized by both high iMPCs and high openness to trade, which are particularly responsive to monetary shocks. The fact that we have estimated a positive and significant interaction term in the data, as Figure 5 has shown, lends support to this key mechanism of our model and serves as an important validation test of the theory.

In Appendix C, Figure C.13, we show that our empirical results are robust to the specific choice of the baseline group (whether it is the top 50, 25, or 10% of the MPC and non-tradable employment distribution), as well as to using narratively identified monetary shocks, as in Romer and Romer (2000), Miranda-Agrippino and Rey (2020), and standard high frequency monetary shocks, as in Gertler and Karadi (2015).

To sum up, in this section, we have documented that US monetary policy shocks in-

Table 1: Model calibration of county-invariant parameters

Parameter	Description	Value	Comment
	Externally Calibrated		
σ	Inverse IES	1	Standard
φ	Frisch Elasticity	1	Chetty et al. (2011)
ν	Cons. elasticity of subs. btw sectors	1.5	Hazell et al. (2022)
θ	Elasticity of subs. btw tradables	4	Hazell et al. (2022)
η	Labor elasticity of subs. btw sectors	0.42	Berger et al. (2022)
μ	Union market power	21	Schmitt-Grohé and Uribe (2005)
ρ_e	Pers. of log-productivity process	0.966	McKay et al. (2016)
σ_e	Std. of log-productivity process	0.504	McKay et al. (2016)
	Internally Calibrated		
$\underline{b}$	Borrowing limit	-1.116	Target $r = 2\%$ (p.a.)

Note: Parametrization of county-invariant parameters in the model.

duce heterogeneous effects across counties. This heterogeneity in county-level responses is systematically explained by two theory-consistent regional characteristics: local MPCs and non-tradable employment shares. Counties with either a high local MPC or a high non-tradable share are significantly more responsive to monetary shocks. Furthermore, we find evidence of a complementarity between MPCs and non-tradable employment in shaping the local response to monetary shocks, consistently with our theory.

4 Calibration

In this section, we turn to the third and final part of our contribution: a quantitative analysis. To begin, we calibrate the model to our empirical estimates of regional MPCs and the share of non-tradable employment. To do so, we propose a methodology to efficiently calibrate the steady state of our model comprising thousands of heterogeneous agent economies linked in general equilibrium. Finally, we describe the characteristics of the joint spatial distribution in the steady state of our calibrated model and introduce three counties that are going to serve as narrative case studies for the rest of our analysis: Manhattan (New York), Luce County (Michigan), and Putnam County (West Virginia).

4.1 Standard parameters

Some of our parameters are standard and county-invariant. Table 1 summarizes them. The top panel describes the parameters that are externally calibrated. As is standard in the literature, we set both the inverse elasticity of intertemporal substitution, σ, and the Frisch elasticity, φ, to 1. We follow Hazell et al. (2022) to calibrate trade elasticities: we set the consumption elasticity of substitution between tradable and non-tradable goods to 1.5 and the elasticity of substitution between tradable varieties to 4. For the elasticity of labor supply between sectors, η, we rely on estimates from Berger et al. (2022) and set it to 0.42. We set unions' labor market power to a standard value of $\mu = 21$ (Schmitt-Grohé and Uribe, 2005). Finally, we follow McKay et al. (2016) for calibrating the income process.[25]

The bottom panel of Table 1 shows the only aggregate parameter that is calibrated internally: the borrowing limit, $\underline{b}$. As we discuss more in detail in the next section, we set it to a value of -1.161 in order to target an annual real interest rate of 2% in the steady state.

4.2 Methodology

We discipline the regional structure of our economy based on two empirical targets: our regional estimates of MPCs and the share of non-tradable employment.[26] To match these targets, we allow two sets of parameters to be county-specific: the discount factor, β_j, and the preference parameter governing non-tradable labor supply, α_j.[27] We choose these parameters to ensure that the first entry of the iMPC Jacobian for county j, $\boldsymbol{M}_j[0,0]$, matches our empirical estimate for that county's MPC and that the model-implied share of non-tradable labor income, ρ_j, equals the corresponding value of $\tilde{\rho}_j$ obtained in our data. We face a considerable numerical challenge of calibrating an economy consisting of more than 3,000 heterogeneous-agent economies linked through the asset market and the market for tradable goods, all in general equilibrium. Building on our theoretical concepts of national and regional equilibrium, we develop a novel two-step methodology to address this challenge. First, we calibrate the full national equilibrium for an economy consisting of 20 counties, carefully chosen to be representative of the full population of US counties. Having solved for national steady-state prices and quantities, we then leverage the atomicity of counties to individually calibrate a regional equilibrium for all the 3,140 counties.

Step 1: national equilibrium calibration. In the first step, we extract 20 representative counties from the data. More specifically, we take 400 million independent draws of 20 counties and select the draw that minimizes the distance between five key moments in the

[25] The income process is discretized with 20 gridpoints using Tauchen's method (Tauchen, 1986).

[26] In particular, we calibrate the model to our regional estimates of MPCs and non-tradable employment for the US economy in 2016.

[27] We set $\omega_j = \alpha_j$ in every county. This assumption enables us to match the non-tradable share exactly in every county.

sample of 20 regions and those in the full U.S. population of counties. We choose the target moments that play a role in shaping our national Keynesian cross (27): the mean and standard deviation of MPCs and non-tradable employment shares, together with the covariance between these variables.[28] Armed with a representative sample of twenty counties, we then calibrate the parameters $\{(\beta_j, \alpha_j)\}_{j=1}^{20}$ to match the national steady state levels of $\{(MPC_j, \rho_j)\}_{j=1}^{20}$ for each representative county.[29] Thus, the first step of our calibration procedure yields the calibrated steady state of a national economy comprising twenty interlinked representative incomplete market economies which is both consistent with the data and computationally feasible.[30]

Step 2: regional equilibrium calibration. Having calibrated the nationwide general equilibrium steady state, we then move to the second step of our procedure. In this step, we leverage the notion of regional equilibrium and calibrate each of our 3,140 counties individually as small open economies. Since each region is atomistic, individual counties do not affect nationwide equilibrium objects. When calibrating individual counties, we can thus take the steady-state interest rate obtained from the national calibration—which we target to be 2% p.a.—as well as the aggregate steady-state national demand for tradables as given. We then sequentially calibrate α_j and β_j for each county to match ρ_j and MPC_j while ensuring local market clearing at the county level, as per Definition 1. This procedure enables us to calibrate all the counties in our sample while ensuring that both national and local markets clear. We hence have all the objects needed to obtain each county's response to shocks: ρ_j, $\boldsymbol{M}_j$, and $\boldsymbol{M}_j^r$.

4.3 The regional distribution of MPCs and openness to trade

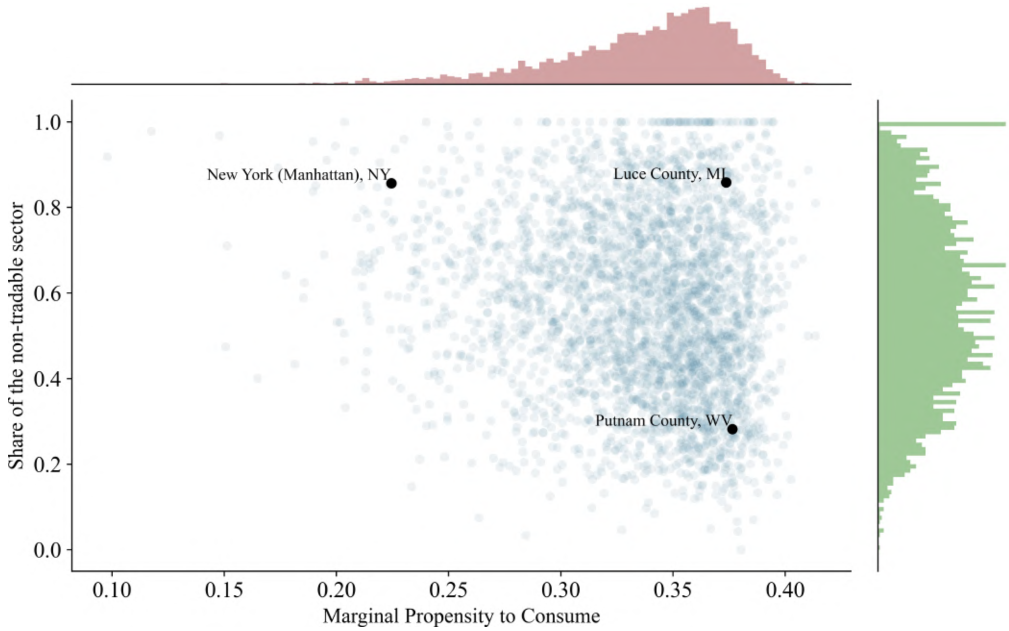
We now describe in more detail the characteristics of our calibrated national steady state. As shown in Table C.1, the calibrated model closely replicates the key moments of the joint distribution of the marginal propensity to consume, MPC_j, and the non-tradable share, ρ_j, observed in the data. Overall, Table C.1 shows how the first step of our calibration strategy successfully matches the empirical regional distribution observed in the US economy. The average MPC is around 30%, which is in the range of standard estimates; the average non-tradable labor income share is 60%, in the ballpark of the consumption share of non-tradable goods in Hazell et al. (2022) (set to 66%). We also match the cross-county standard

[28] Our target moments are all population-weighted.

[29] We jointly calibrate the county-invariant borrowing limit to target a steady state interest rate of 2% p.a.

[30] We list the values we choose for $\{(\beta_j, \alpha_j)\}_{j=1}^{20}$, together with our empirical targets for $\{(MPC_j, \rho_j)\}_{j=1}^{20}$ in Table C.2 in Appendix C.

Figure 6: Joint regional distribution of MPCs and non-tradable employment



Note: Each of the 3,140 dots represents a separate county in the data as well as the steady state of the calibrated model. The histograms plot the marginal distributions of MPCs and non-tradable shares. Since these two statistics are matched exactly in the model, the graph represents simultaneously the empirical and the model-based steady-state distribution.

deviation in MPCs and non-tradable shares, around 4% and 15% respectively, as well as the negative covariance between the two.

Figure 6 plots the joint steady-state distribution of MPCs and non-tradable employment for the fully calibrated set of 3,140 counties. Each dot in the cloud represents the MPC and non-tradable employment share of the model counterpart of a real-world county in our dataset.[31] Thus, each dot in Figure 6 corresponds to a calibrated heterogeneous-agent economy with its own steady-state prices and distribution of household wealth. We also report the marginal distributions of MPCs and non-tradable employment across US counties in the top and right histograms, respectively. Because we calibrate our model to exactly match the measured MPC and non-tradable employment shares, these empirical distributions are the same as those in our model.

Narrative case studies. We now present three counties that are going to serve as narrative model-based case studies for the rest of our analysis: Manhattan (New York), Luce County (Michigan), Putnam County (West Virginia). These counties are marked with black dots and labelled accordingly in Figure 6. First, we choose New York for its salience and because it is a typical example of a region with low MPCs and high non-tradable

[31] Figure C.10 presents the binned scatterplot version of this graph.

employment shares. Manhattan's economy is heavily dominated by services and finance, sectors that are generally non-tradable. Moreover, residents of Manhattan tend to have high income and wealth, which translates into low estimated MPCs in our empirical analysis. In contrast, we pick Putnam County, West Virginia to exemplify high MPC, low non-tradable employment regions. Its economy relies heavily on mining and manufacturing, placing it in the bottom 10% of the non-tradable employment share across US counties. At the same time, Putnam County ranks in the top 20% of our estimated MPC distribution, reflecting the historically low income and wealth levels characterizing the Appalachian region. Finally, we choose Luce County, Michigan as our third and last example county. We pick this region as it has the same trade openness as New York as well as the same MPC as Putnam, thus making it particularly well suited to illustrate the different channels at play in our model. To further illustrate the numerical properties of our model, Figure C.1 in Appendix C plots steady-state Jacobians $\boldsymbol{M}_j$ and $\boldsymbol{M}_j^r$ associated with these counties, together with their steady-state distributions and policy functions.

In summary, we propose a two-step approach to bring the model to the empirical regional estimates of MPCs and non-tradable employment. We calibrate the model to match the first- and second-order moments of the regional distribution of MPCs and non-tradable shares observed in the data. Our calibrated model delivers realistic heterogeneity in local MPCs and non-tradable employment for the full set of 3,140 U.S. counties. This joint distribution is at the centre of our quantitative results, to which we now turn.

5 Quantitative results

This section presents the main quantitative results from our calibrated model. We begin by inspecting county-level responses. First, our framework successfully replicates the empirical patterns documented in Section 3. We also provide a decomposition of the county-level response to monetary shocks into three separately additive channels and show how the nature of the monetary transmission mechanism varies across regions. We move on to discuss the distributional effects of monetary policy across space. Our two dimensions of regional heterogeneity generate sizeable dispersion in the regional response to common monetary shocks. Moreover, our framework predicts that expansionary monetary policy reduces regional consumption inequality. Finally, we discuss the implications of regional heterogeneity for the aggregate transmission of monetary policy. To this end, we introduce and discuss additional regional data for Italy. We calibrate our model to this Italian data as well and demonstrate how differences in the American and Italian geographies have potentially large implications for the aggregate transmission of monetary policy.

5.1 Matching regional impulse responses

We begin by showing that our model is able to reproduce the empirical impulse responses documented in Figure 5. From our calibrated model economy we can compute nationwide responses to a transitory monetary policy shock as a fixed point between the regional and national Keynesian crosses (20) and (27), for a given impulse $\hat{\boldsymbol{r}}$. We set the quarterly persistence of monetary shocks to 0.61, in line with Kaplan et al. (2018). This yields the nationwide labor response $\hat{\boldsymbol{\ell}}$, as well as the national response of tradable consumption $\hat{\boldsymbol{c}}^T$. We can then leverage our regional Keynesian cross equation (20), together with the calibrated triad of $(\boldsymbol{M}_j, \boldsymbol{M}_j^r, \rho_j)$ for all 3,140 counties, to obtain model-implied responses of every single county in the economy. Finally, we run the same panel local projection on the model-simulated output as we did for the US data in (31):

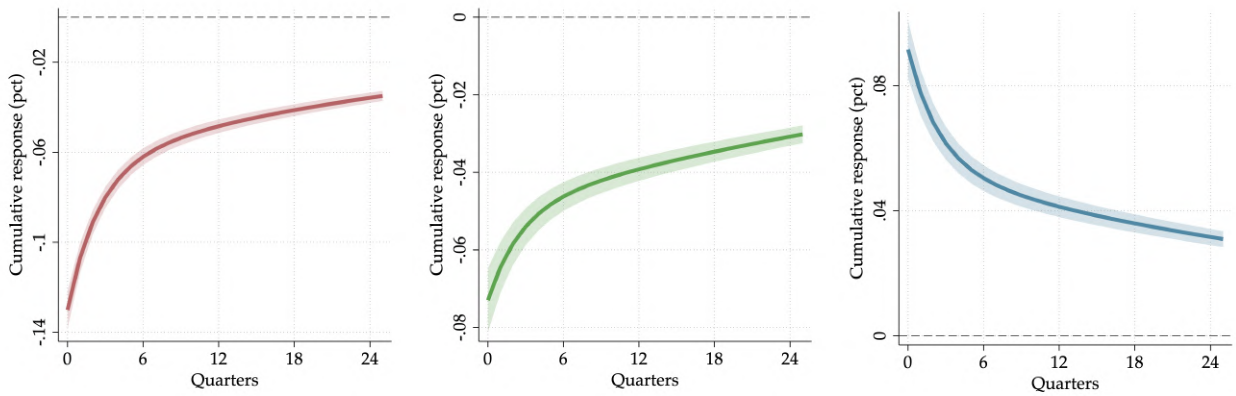
$$\hat{\ell}_{jt+h} = \lambda_h + \gamma_h^{\rho} D_j^{\rho} + \gamma_h^{M} D_j^{M} + \gamma_h^{\rho\times M} D_j^{\rho} \times D_j^{M} + u_{jht} \tag{32}$$

where $\hat{\ell}_{jt+h}$ is the model-simulated log-deviation of employment from steady state in county j, h quarters after the shock, λ_h is a constant, and D_j^{ρ} and D_j^{M} are defined in the same way as in our empirical specification in Section 3. In this exercise, we consider a monetary shock normalized to increase national employment by 1% on impact in order to facilitate a cleaner comparison. Figure 7 plots the estimated coefficients from equation (32), and is the model analog to Figure 5.[32] As before, the baseline group of counties is represented by those that are in the top quintile of both the MPC and the non-tradable employment distribution. Because we are considering an expansionary monetary shock, the average employment response among this group of counties is positive.

The left panel of Figure 7 shows that, relative to this baseline group, counties with low MPCs but high non-tradable employment tend to be *less* responsive to interest rate changes. This finding may appear surprising in light of the results in Werning (2015), who shows how, in the general context of HANK economies, the effect of MPCs on the general equilibrium response to interest rate changes is ambiguous. This ambiguity is not present in our framework at the county level. In our model, the regional equilibrium labor response is shaped by two simultaneous impulses to the local economy. First, the interest rate shock and, second, the change in national demand for tradable goods. While high MPCs tend to amplify indirect general equilibrium effects, they also tend to be associated with a smaller direct consumption response to interest rate changes—hence the ambiguity. However, in our setting, the direct response of the local economy to the tradable demand

[32] In Figure B.1 in Appendix B we repeat the same exercise in the version of our model featuring sticky wages and show that the results are unaffected.

Figure 7: Model-based local projection estimates



Note: model-based impulse response functions capturing the coefficients in (32) in response to an expansionary monetary policy shock that raises average employment by 1% on impact with quarterly persistence of 0.61. Shaded areas represent 95% confidence intervals. The y-axis is in units of the cumulative percentage change in employment.

shock does not depend on county-level MPCs. Given this, the regional employment response is increasing with county-level MPCs, as national demand acts as an additional impulse feeding the regional Keynesian multiplier.

Similarly, the middle panel of Figure 7 shows how counties with low non-tradable employment but high MPCs exhibit a milder employment response to a monetary expansion. This is in line with our empirical estimates and with the intuition that a smaller ρ dampens the regional Keynesian multiplier, as discussed in the context of Figure 1.

Finally, the right panel of Figure 7 presents results for $\hat{\gamma}_h^{\rho,M}$. As discussed before, this coefficient has the interpretation of a cross-derivative, capturing how MPCs (non-tradable employment) mediate the effect of non-tradable employment (MPCs) on the local response. Hence, the positive and significant estimate for $\hat{\gamma}_h^{\rho,M}$ indicates that the dampening effect of low MPCs (non-tradable employment) on the county-level response to a monetary shock is weaker for counties with low non-tradable employment (MPCs). In other words, counties with both high local MPC and a large non-tradable sector are more sensitive to monetary impulses. This result is consistent with our empirical findings and is in line with the intuition behind the regional Keynesian multiplier $\boldsymbol{\mathcal{M}}_j \equiv \left(\boldsymbol{I} - \rho_j \boldsymbol{M}_j\right)^{-1}$, discussed in Figure 1. As the share of non-tradable employment and MPCs enter this expression multiplicatively, they complement each other in shaping the magnitude of the regional multiplier $\boldsymbol{\mathcal{M}}_j$. This complementarity is precisely what the positive estimated $\hat{\gamma}_h^{\rho,M}$ reflects and is going to be crucial in generating aggregate implications from regional distributions in the remainder of this section.

5.2 Channels of monetary policy transmission across space

Having established that the calibrated model can match the empirical regularities in the regional responses to monetary policy shocks, we now decompose the various transmission channels of monetary policy that are at play at the regional level. An important contribution of the canonical HANK literature has been to shed light on the role household heterogeneity plays in the transmission mechanism of economic policy in standard single-region models (Kaplan et al., 2018, Auclert et al., 2024a). We now show how in our multi-region framework the relevance of the different transmission channel varies considerably across space.

Case studies. We begin by zooming in on the three case-study examples of Manhattan (NY), Luce County (MI) and Putnam County (WV), which we first introduced in Section 4.3. Figure 8 plots employment responses to the same expansionary monetary impulse for these three counties in separate panels.[33] The central panel of Figure 8 plots the response of the baseline county, Luce County (MI), which has a high MPC and a high non-tradable share. The left panel is New York (NY), which has a low local MPC and a high non-tradable share. Finally the right panel is Putnam County (WV) which has a high MPC and low trade openness.

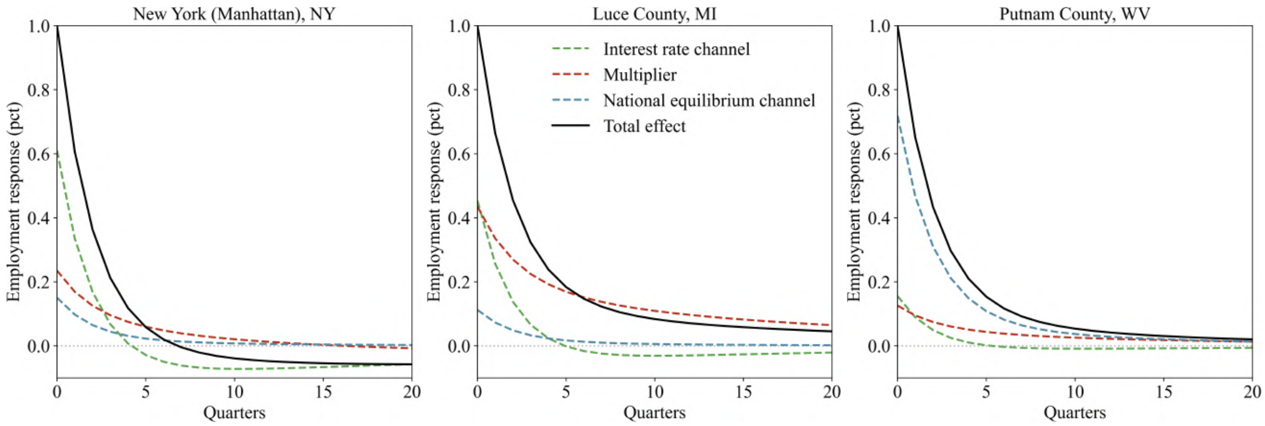
In each panel of the Figure, we present the total response (normalized to 1% on impact) and decompose it into three separately-additive channels. The regional Keynesian cross equation (20) provides an exact decomposition of the employment response at the county level into a standard interest rate channel, the regional multiplier, and the national equilibrium channel:

$$\widehat{\boldsymbol{\ell}}_j = \underbrace{\rho_j \boldsymbol{M}_j^r \widehat{\boldsymbol{r}}}_{\text{Interest rate channel}} + \underbrace{\rho_j \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}_j}_{\text{Regional multiplier}} + \underbrace{(1-\rho_j)\widehat{\boldsymbol{c}}^T}_{\text{National equilibrium channel}} \tag{33}$$

The main takeaway from Figure 8 is that the local transmission of monetary policy for these three counties occurs via very different channels. This takes place because of ex-ante cross-county heterogeneity in MPCs and trade openness. Consider first that the only difference in the responses of New York and Luce counties comes from their different MPCs, as the non-tradable share is the same across the two. From the perspective of an individual county—in a regional equilibrium—the monetary shock induces two impulses that are exogenous to the local economy: the direct interest rate change as well as the national demand for the tradable good. The latter channel is identical across the two counties since

[33] Figure C.2 in Appendix C.1 plots the same IRFs with a longer truncation horizon.

Figure 8: Case studies of heterogeneous responses to monetary policy



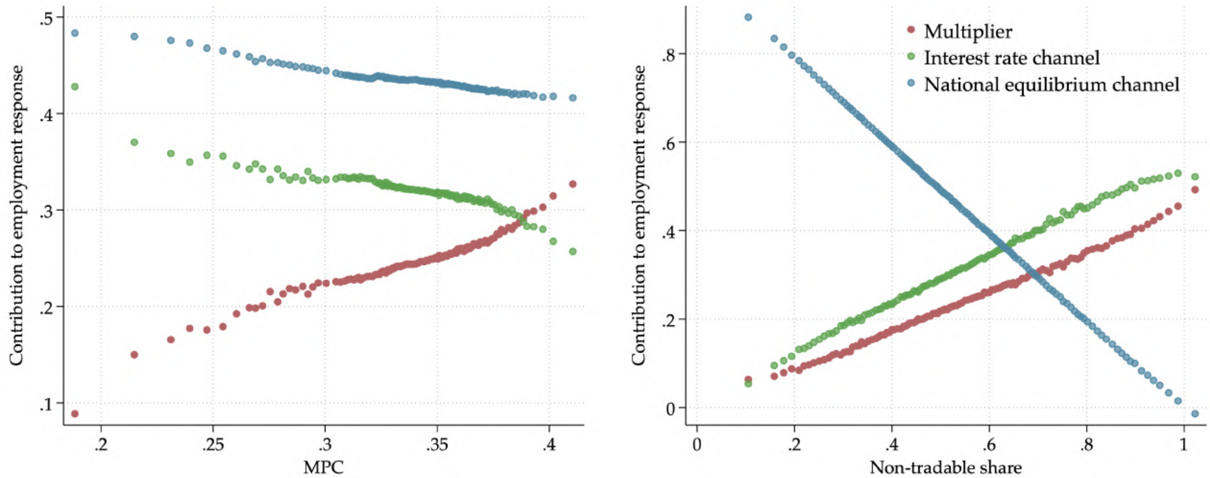
Note: decomposition of the employment response to a monetary shock that increases employment by 1% in each county with quarterly persistence of 0.61.

both the non-tradable employment share and the sequence of national tradable demand are common. However, because MPCs are higher in Luce County, the multiplier channel is much larger than in New York (40% of on-impact response in Luce County against 25% only in New York), while the interest rate channel is smaller. On the other hand, the response of the high-MPC, high-openness Putnam County is primarily driven by the national response, while the effect of the interest rate channel and especially the multiplier channel are dampened by the low share of the non-tradable sector.

Systematic decomposition. Motivated by the three case studies above, we now investigate more systematically the contribution of the different transmission channels of monetary policy across space. Figure 9 shows how the contribution of the interest rate, regional multiplier, and national equilibrium channels varies across the joint distribution of regional MPCs and non-tradable employment in our full sample of counties. For each county, we leverage the regional Keynesian cross decomposition (33) to compute the share of the on-impact employment response coming from each channel. In the left panel, we group counties into 120 bins based on their local MPC. For each bin, we then plot the average contribution to the employment response—residualized by county-level non-tradable shares—coming from each channel. In the right panel, we repeat the same procedure, this time grouping counties based on their non-tradable employment share and residualizing by county-level MPCs.

Focusing on the left panel of Figure 9, our first observation is that the contribution of the regional Keynesian multiplier is increasing in MPCs, which is consistent with our earlier discussions. Conversely, the contribution of the interest rate channel is decreasing

Figure 9: Channels of monetary policy transmission



Note: binned scatterplots of the county-level contributions to the on-impact response of employment to a monetary shock. Non-tradable shares are controlled for in the left panel, and MPCs are controlled for in the right panel.

with MPCs. This is a standard feature of the interest rate channel in the canonical HANK model (Kaplan et al., 2018). The contribution of the national demand channel is slightly decreasing in MPCs because the level of the responses is increasing in MPCs while the national equilibrium channel is constant over MPCs, as discussed at the beginning of this section.

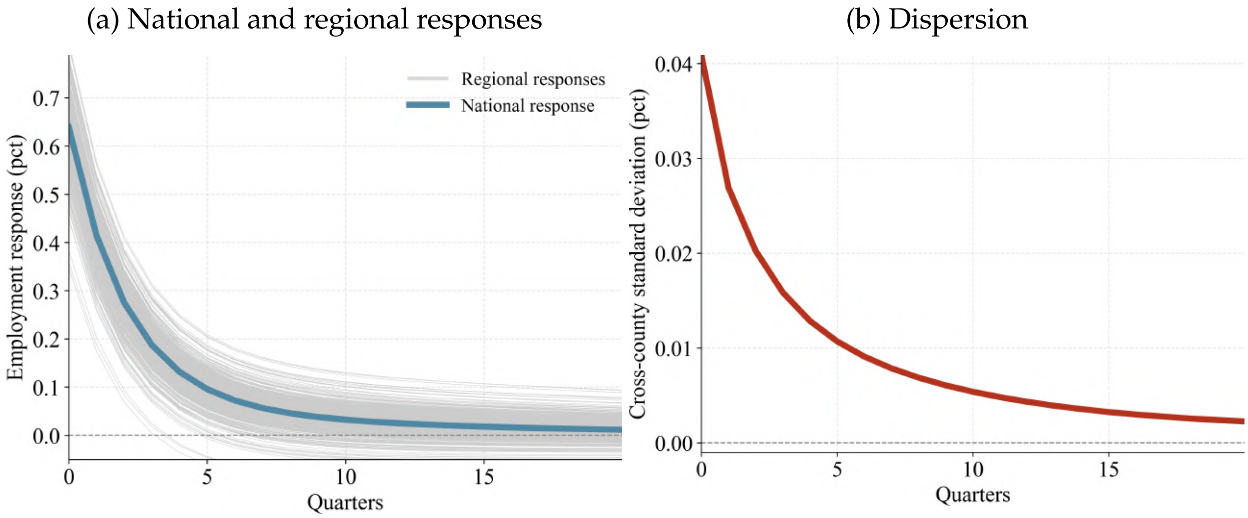
Moving on to the right panel of Figure 9, we see that most of the variation in the total response is driven by the shift away from the national and towards the regional equilibrium effects as the non-tradable share increases. As such, note that as local non-tradable employment increases, the share of the national demand channel falls linearly from unity to zero and contribution from both the interest rate and the multiplier channels increases. Moreover, we also see how indirect effects gain importance non-linearly as the share of non-tradable employment, ρ_j, goes to one.

Overall, in this section, we have shown how regional differences in MPCs and non-tradable employment give rise not only to heterogeneity in the level of the local response to monetary policy shocks, but also in the mechanism via which monetary policy affects local economic activity.

5.3 Distributional effects of monetary policy

Having analyzed the heterogeneous transmission of monetary policy across regions, in this section, we now turn to study how this heterogeneity shapes the distributional ef-

Figure 10: National and regional effects of monetary policy



Note: the left panel plots the regional and national employment responses to a monetary shock that decreases r_t by 1 p.p. (annualized) on impact with quarterly persistence of 0.61. The right panel plots the cross-county standard deviation of percentage employment deviations.

fects of monetary policy across space. The left panel of Figure 10 plots the nationwide employment response to an expansionary monetary policy shock, together with the full set of county-level responses.[34] The shock raises national employment by around 0.6 per cent on impact, with a half-life of around three quarters. Importantly, the two dimensions of regional heterogeneity present in our model generate sizeable dispersion in the local response to the shock. Employment increases by as much as 0.8% in the most responsive counties and as little as 0.35% in the least responsive ones.

To more formally quantify the regional dispersion in the response to monetary shocks, the right panel of Figure 10 plots the IRF of the cross-sectional standard deviation of county-level employment responses. This measure captures the dispersion at different horizons in the county-level IRFs represented by the grey shaded area in the left panel of Figure 10. As already mentioned, a monetary easing increases the dispersion in regional employment deviations from steady state.[35] What the right panel of Figure 10 also shows is that the effect on regional dispersion is more persistent than the effect on national employment itself. While the half-life of the national response is three quarters, the dispersion in regional responses halves after as much as seven quarters.

How does the degree of regional differences in the response to monetary shocks generated by our model compare to what we estimated in the data in Section 3? To address this question, we need to account for the fact that empirical county-level responses con-

[34] Figure C.3 in Appendix C.1 plots the same IRFs with a longer truncation horizon.

[35] The effect does not depend on the sign of the shock. In other words, a monetary tightening would increase dispersion in county-level employment responses as well.

flate "true" regional heterogeneity together with the dispersion that is purely attributable to estimation error. To account for this, we follow the procedure used in, among others, Chetty et al. (2014) and Kleven (2024). Under the assumption that estimation noise is uncorrelated with the estimated coefficient itself, one can decompose total variance in the estimated $\hat{\beta}_j$—in an additively-separate form—into the true variance and the variance that is due to estimation noise. We proxy the latter with the sum of the squared standard errors, and retrieve the true variance residually. Applying this procedure, we find that the resulting dispersion in the regional responses to a monetary shock generated by our model is 20% of what we estimate in the data.[36] Recall that our model features just two dimensions of regional heterogeneity and abstracts from many others that are also present in the data—such as housing markets or banking access. Despite that, regional differences in MPCs and non-tradable employment appear to be first-order determinants of the monetary policy transmission mechanism across space.

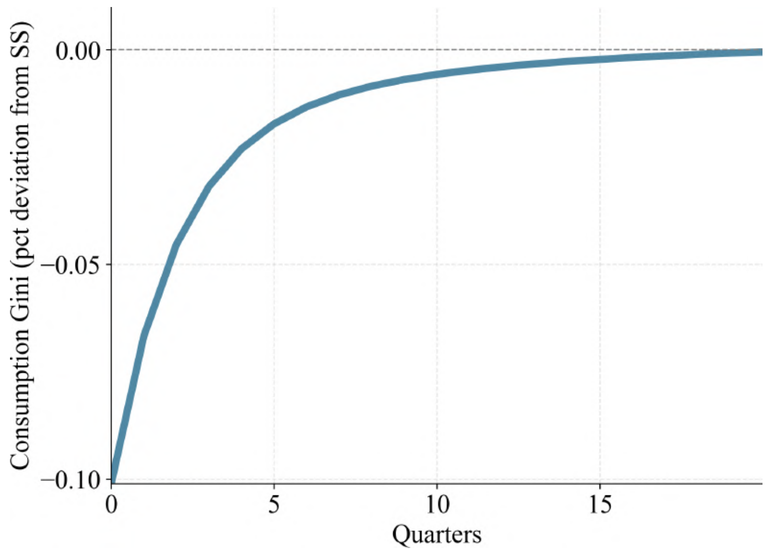
The analysis in Figure 10 only captures the dispersion in county-level responses, and thus speaks to the extent to which monetary policy has different effects across space. However, it is not in itself informative about the effect of monetary policy on the level of regional inequality. This is the focus of the next paragraph.

Consumption inequality. Motivated by the previous discussion, we now explore the effects of monetary policy on regional inequality. Specifically, we focus on consumption inequality and construct the Gini index for county-level consumption across our 3,140 counties. In Figure 11, we plot the response of this Gini index (in log-deviations from its steady state value) to the same monetary shock considered before. Following the same expansionary monetary shock as above the consumption Gini coefficient decreases by around .17%, roughly one quarter of the magnitude of the national employment response. Thus, expansionary monetary policy reduces regional inequality.

The intuition behind this result is the following: high-MPC regions tend to have low steady-state consumption, but they also happen to be more responsive to interest rate changes Thus, monetary easing particularly stimulates consumption growth in areas where consumption levels are lower initially. As a result, this curbs regional inequality as low-consumption counties catch up to the high-consumption ones. Clearly, the opposite is true in the case of a monetary tightening. Thus, our quantitative framework predicts that restrictive monetary policy exacerbates regional inequality while expansionary policy dampens it. Consistently with our previous finding, the effect of monetary policy

[36] For cleaner comparability, we normalize the shock so that the national employment response predicted by our model is the same as what we estimate in the data. We plot the full distribution of regional responses to monetary policy in the model and in our data in Figure D.1 of Appendix C.

Figure 11: Consumption inequality and monetary policy



Note: the figure shows the response of the regional Gini coefficient for consumption to the same shock as in Figure 10. The y-axis is in percentage deviations from steady state.

on regional inequality is more persistent than its effect on aggregate employment, as the half-life of the regional Gini response is about five quarters.

In summary, in this section, we have analyzed the predictions of our model for the distributional effects of monetary policy across space. First, we show that county-level employment responses to the same monetary shock vary substantially across regions. This heterogeneity in the local response implies that monetary policy affects regional inequality. Moreover, the underlying regional disparities persist long after the shock, further highlighting the importance of considering geographic heterogeneity when assessing the overall impacts of monetary policy decisions.

5.4 Aggregate implications of regional heterogeneity: the U.S. and Italy

Beyond speaking to heterogeneity across regions, our model also has predictions for the aggregate, nationwide transmission of monetary policy. In particular, the national Keynesian Cross (27) emphasized how the aggregate effects of monetary shocks are shaped by the joint regional distribution of MPCs and non-tradable employment. In this section, we show that this mechanism can be quantitatively important. This means that two economies that are identical from a national perspective, i.e., share the same steady-state values for nationwide objects, can exhibit very different aggregate responses to monetary shocks if their underlying regional distributions differ. To illustrate this, we examine quantitatively the implications of regional heterogeneity in our model by comparing two baseline economies with distinct joint spatial distributions of MPCs and trade openness: the United States and Italy. We choose Italy because it has readily available data on MPCs at the regional level and because its regional structure differs from that of the United States, as we will describe momentarily.

Italian data. We now introduce and discuss microeconomic data for Italy. We use the 20 Italian NUTS2 regions as our geographical unit of interest. To construct a regional measure of non-tradable employment we use sectoral regional employment statistics from the Italian National Institute of Statistics.[37] To measure regional MPCs, we rely on the widely-used Italian Survey of Household Income and Wealth (SHIW), which is run biannually by the Bank of Italy.[38] The SHIW is a representative survey of Italian households, featuring a rotating panel component. It includes detailed disaggregated data on households' income, assets, and liabilities, and also provides information on consumption and saving behavior, as well as demographic characteristics.

Following Auclert (2019), we focus on the 2010 wave of the survey, when respondents were asked to self-report the portion of a hypothetical windfall gain that they would spend immediately (Auclert, 2019, Jappelli and Pistaferri, 2020).[39] The advantage of this survey is that it contains direct information about households' MPCs and that its sample size is large enough to include a sufficient amount of observations within each of the twenty Italian regions we consider. More specifically, the 2010 wave of the survey covers 7,951 respondents, leaving us with an average of 397 MPC observations per region. Hence, we construct our regional MPC measure as the average of the MPCs reported by all respondents within a region, weighted by survey weights.[40]

Figure C.4 in Appendix C plots the resulting geographic distributions of MPCs and the non-tradable shares across Italian regions. When calibrating our baseline multi-region HANK model to the Italian economy, we use the same externally calibrated parameters as in Table 1, and follow the same procedure described in Section 4.2 for setting the values of the remaining parameters. However, we do not extract representative counties as in our

[37]These data have a coarser industry classification than the CBP we use for the US. For this reason, we classify as non-tradable the "Retail trade, accommodation and restaurants" sector and as tradable the "Industry, excluding construction" sector. We then construct the empirical $\tilde{\rho}_{jt}$ following the same procedure as in Section 3.2.

[38]This survey has been used extensively in the literature. See, for example, Jappelli and Pistaferri (2014, 2020), Auclert (2019).

[39]The size of this hypothetical windfall gain is equal to the household's monthly income.

[40]A well-known issue with the SHIW is that the MPC question does not specify the time frame over which for the MPC, as discussed in Jappelli and Pistaferri (2014) and Auclert (2019). In our baseline calibration, we interpret the MPCs as referring to a quarterly time frame. However, our results are robust to a recalibration where the reported MPCs are interpreted as referring to an annual frequency.

two-step calibration procedure, as our Italian sample only includes twenty regions, which we thus target directly.

Comparing Italian and American geographies. Table C.4 in Appendix C compares regional heterogeneity for the Italian and U.S. geographies. The two economies share a similar aggregate sectoral composition, with the average share of non-tradable employment being 60% in the U.S. and 55% in Italy. The average MPC for the Italian economy is larger than in the U.S., in line with what has been previously documented by, for example, Jappelli and Pistaferri (2014) and Auclert (2019).

More relevant for our purposes is the fact that the Italian economy features much more regional dispersion in MPCs than the American one. The cross-sectional standard deviation of MPCs is only 3.9% in the U.S., while it is 13.9% in Italy. Since our theory predicts that covariances are the key factors in determining how regional heterogeneity influences the aggregate transmission of shocks, this will play an important role in shaping our results. Cross-sectional dispersion in trade openness is, instead, comparable across the two countries, with the standard deviation of the non-tradable share of employment being 16.1% in the U.S. and 11.3% in Italy.

The last relevant difference across the two countries is in the covariance between MPCs and non-tradable employment across space. The Italian economy features a strong *positive* correlation of 34.2% between MPCs and non-tradable shares across regions. On the other hand, this correlation in the U.S. economy is of similar magnitude but of opposite sign, -31.4%.

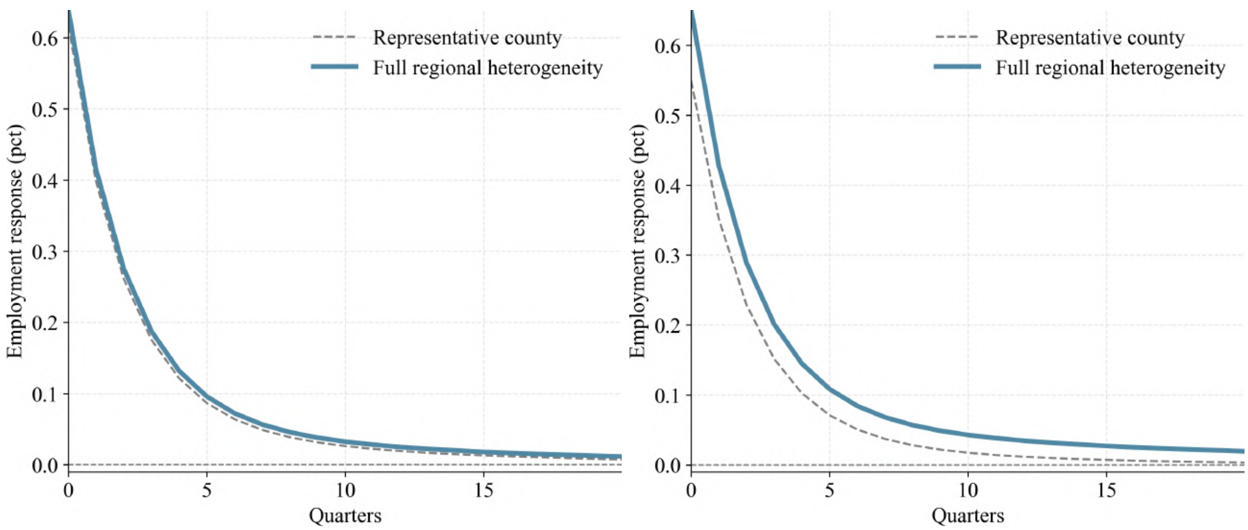
Amplification. We now quantify how the joint spatial distribution of MPCs and openness to trade shapes the nationwide macroeconomic response to monetary policy shocks. To do so, we construct counterfactual versions of the U.S. and Italian economies, recalibrated to eliminate regional heterogeneity while preserving the average nationwide MPC and steady-state interest rate from the baseline heterogeneous economies.[41] We then compute aggregate employment responses to the same monetary shock in the American and Italian economies with full regional heterogeneity and in their counterfactual representative-county versions.

Panels (a) and (b) of Figure 12 present the results of this exercise for the U.S. and Italian economies, respectively.[42] Qualitatively, in both countries we find that regional

[41] In particular, we change the borrowing limit $\underline{b}$ to target the steady state interest rate and set the discount factor β to target the first entry of the iMPC Jacobian.

[42] Figure C.5 in Appendix C plots the cumulative response.

Figure 12: Regional heterogeneity and aggregate amplification



(a) United States (b) Italy

Note: the figure plots employment responses to a monetary shock that decreases r_t by 1 p.p. (annualized) on impact with quarterly persistence of 0.61, in the full model vs. the representative-county case. The representative county stands for the response of the re-calibrated economy with no regional heterogeneity while the full regional heterogeneity response refers to the baseline heterogeneous-county model from the main text.

heterogeneity amplifies the aggregate impact of monetary policy. However, regional heterogeneity turns out to be quantitatively unimportant for the aggregate transmission of monetary policy in the U.S., amplifying the response by just 3% on impact and 13% cumulatively. In the case of Italy, instead, regional heterogeneity amplifies the aggregate response by as much as 18% on impact and 59% cumulatively.[43] This result suggests that the extent to which regional distributions matter for aggregate dynamics depends on the nature of spatial heterogeneity itself. We now explain the intuition behind this key quantitative result.

Figure C.7 in Appendix C leverages the national Keynesian cross in Proposition 3 to decompose the source of this amplification into direct and indirect effects. This exercise shows that it is the indirect effect channel, i.e. the regional Keynesian multiplier, that is the main driver through which regional differences amplify the national response. While the magnitude of the direct effect under regional heterogeneity is only mildly larger than in the representative-region benchmark, the introduction of spatial heterogeneity increases the multiplier channel by as much as 24% in the case of Italy. Recall that because the multiplier is non-linear and convex in both MPCs and non-tradable employment, greater dispersion in either factor generates stronger amplification. Furthermore, recall

[43]Figure C.6 in Appendix C shows that this result is robust to the MPC calibration. In particular, when we recalibrate the Italian economy interpreting our regional MPCs as referring to an annual frequency, regional heterogeneity still delivers amplification of 10% on impact and 43% cumulatively.

that MPCs and non-tradable employment act as complements in shaping the regional Keynesian multiplier. A positive spatial correlation between the two, hence, provides an additional source of amplification from regional differences. Consequently, in the U.S., the heterogeneity arising from dispersion in MPCs and non-tradable employment slightly outweighs the dampening effect stemming from their negative correlation. In contrast, Italy exhibits both larger dispersion in MPCs and a positive correlation between MPCs and non-tradable employment—–two forces that jointly deliver substantial amplification.

Finally, in Figure B.2, we show that our results are both qualitatively and quantitatively robust to the assumption of rigid wages. Under sticky wages, the aggregate amplification brought about by regional heterogeneity in the Italian economy remains virtually unchanged, at 19% on impact and 38% in cumulative terms.

Taking stock, this section has demonstrated that geographic heterogeneity in MPCs and trade openness can have quantitatively significant effects on the transmission of monetary policy in a realistically calibrated model. Moreover, we have highlighted how the nature of spatial heterogeneity shapes the extent to which regional distributions influence aggregate dynamics. This suggests that the role of regional heterogeneity in the national transmission of shocks can vary across geographies. Finally, we have outlined the key features of the geographical distribution that policymakers must consider and keep track of when assessing the importance of regional heterogeneity for the aggregate effectiveness of monetary policy.

6 Conclusion

In this paper, we examine the role of regional heterogeneity in the aggregate transmission of monetary policy and its distributional effects across space. We develop a multi-region heterogeneous-agent model with asymmetric counties and show how regional differences in iMPCs and non-tradable employment shares shape both local and national responses to monetary policy shocks. We cast our model in the sequence space and derive the regional Keynesian cross, a closed-form expression for the local employment response to a monetary shock. At the center of this expression is the regional Keynesian multiplier, a non-linear function of iMPCs and trade openness. We then turn to aggregation and derive the national Keynesian cross, which characterizes the national employment response in terms of the joint regional distribution of iMPCs and trade openness. Using U.S. county-level micro-data, we construct a novel measure of regional MPCs to empirically validate our model. Finally, we calibrate our model to the U.S. and Italian geographies and show that spatial heterogeneity can substantially amplify monetary policy shocks.

Looking ahead, the paper opens several avenues for future research. Our framework is portable and can be applied to different contexts. In Bellifemine et al. (2025), we extend it to the context of the euro area to study how differences in legacy public debt across member countries shape the transmission of monetary policy. Future work could explore other important dimensions of regional heterogeneity, such as differences in labor mobility or wage and price rigidities. More broadly, applying our framework to countries with different regional structures could enhance our understanding of how geography influences macroeconomic dynamics around the world.

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Online Appendix

The Regional Keynesian Cross

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A Appendix to Section 2

Here we provide proofs to the Lemmas and Propositions in Section 2.

A.1 Proof of Lemma 1

By the definition of the labor aggregator, we have:

$$\begin{aligned}
\widehat{\boldsymbol{\ell}}_j &= \frac{\eta}{\eta+1}\frac{\eta+1}{\eta}\alpha_j^{-\frac{1}{\eta}}\frac{1}{L_j^{\frac{\eta+1}{\eta}}}(L_j^{NT})^{\frac{1}{\eta}}\widehat{\boldsymbol{\ell}}_j^{NT} + \frac{\eta}{\eta+1}\frac{\eta+1}{\eta}(1-\alpha_j)^{-\frac{1}{\eta}}\frac{1}{L_j^{\frac{\eta+1}{\eta}}}(L_j^T)^{\frac{1}{\eta}}\widehat{\boldsymbol{\ell}}_j^T \\
&= \frac{L_j^{NT}}{L_j}\alpha_j^{-\frac{1}{\eta}}\left(\frac{L_j^{NT}}{L_j}\right)^{\frac{1}{\eta}}\widehat{\boldsymbol{\ell}}_j^{NT} + \frac{L_j^T}{L_j}(1-\alpha_j)^{-\frac{1}{\eta}}\left(\frac{L_j^T}{L_j}\right)^{\frac{1}{\eta}}\widehat{\boldsymbol{\ell}}_j^T \\
&= \frac{L_j^{NT}}{L_j}\frac{W_j^{NT}}{W_j}\widehat{\boldsymbol{\ell}}_j^{NT} + \frac{L_j^T}{L_j}\frac{W_j^T}{W_j}\widehat{\boldsymbol{\ell}}_j^T \\
&= \rho_j\widehat{\boldsymbol{\ell}}_j^{NT} + (1-\rho_j)\widehat{\boldsymbol{\ell}}_j^T
\end{aligned}$$

By market clearing, we have:

$$\widehat{\boldsymbol{\ell}}_j^{NT} = \widehat{\boldsymbol{c}}_j^{NT} = \nu(\widehat{\boldsymbol{p}}_j^{NT} - \widehat{\boldsymbol{p}}_j) + \widehat{\boldsymbol{c}}_j \tag{A.1}$$

Moreover, it holds that $\widehat{\boldsymbol{\ell}}_j^{NT} = \widehat{\boldsymbol{c}}_j$. Finally, given that the demand for county j's tradable good $\widehat{\boldsymbol{c}}^T(j)$ is exogeneous in a regional equilibrium and by imposing market clearing, we have:

$$\widehat{\boldsymbol{\ell}}_j = \rho_j\widehat{\boldsymbol{c}}_j + (1-\rho_j)\widehat{\boldsymbol{c}}^T(j) \tag{A.2}$$

Finally, we have:

$$C_t^T(j) = \int C_{j't}^T(j)dj' = \int \left(\frac{P_{jt}^T}{P_t^T}\right)^{-\theta} C_{j't}^T dj'$$

which implies:

$$\widehat{\boldsymbol{c}}^T(j) = -\psi(\widehat{\boldsymbol{p}_j}^T - \widehat{\boldsymbol{p}_t}^T) + \int \frac{C^T_{j'}}{C^T}\widehat{\boldsymbol{c}}_{j'}dj$$

Now defining $\widehat{\boldsymbol{c}}^T = \int \frac{C^T_{j'}}{C^T}\widehat{\boldsymbol{c}}_{j'}dj$ we have that the variation in demand for the tradable goods produced in county j equals the national variation for tradable goods: $\widehat{\boldsymbol{c}}^T(j) = \widehat{\boldsymbol{c}}^T$.

A.2 Proof of Proposition 1

By the definition of the aggregate consumption function we have:

$$\widehat{\boldsymbol{c}}_j = \boldsymbol{M}_j(\widehat{\boldsymbol{\ell}}_j + \widehat{\boldsymbol{w}}_j - \widehat{\boldsymbol{p}}_j) + \boldsymbol{M}^r_j\widehat{\boldsymbol{r}}_j \tag{A.3}$$

Simplifying:

$$\widehat{\boldsymbol{c}}_j = \boldsymbol{M}_j\widehat{\boldsymbol{\ell}}_j + \boldsymbol{M}^r_j\widehat{\boldsymbol{r}} \tag{A.4}$$

Hence, replacing (A.4) in Lemma 1, we get:

$$\widehat{\boldsymbol{\ell}}_j = \rho_j\left(\boldsymbol{M}_j\widehat{\boldsymbol{\ell}}_j + \boldsymbol{M}^r_j\widehat{\boldsymbol{r}}\right) + (1-\rho_j)\widehat{\boldsymbol{c}}^T \tag{A.5}$$

A.3 Proof of Proposition 3

The percentage change in demand for the tradable good is equal to the weighted change in consumption demand:

$$\widehat{\boldsymbol{c}}^T = \int \frac{C^T_j}{C^T}\widehat{\boldsymbol{c}}_j dj$$

We can define the measure $d\mu_j \equiv \frac{C^T_j}{C^T}dj$ and express the operators with respect to this measure:

$$\widehat{\boldsymbol{c}}^T = \boldsymbol{M}\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + \mathrm{Cov}(\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j)$$

Plugging this back into (20) and integrating we have:

$$\widehat{\boldsymbol{\ell}}_j = \rho_j\left(\boldsymbol{M}_j\widehat{\boldsymbol{\ell}}_j + \boldsymbol{M}^r_j\widehat{\boldsymbol{r}}\right) + (1-\rho_j)\left(\boldsymbol{M}\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + \mathrm{Cov}(\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j)\right)$$

$$
\begin{aligned}
&\implies \widehat{\boldsymbol{\ell}} = \int \rho_j \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}_j d\mu_j + \int \rho_j \boldsymbol{M}^r_j d\mu_j \widehat{\boldsymbol{r}} + (1-\rho)\left(\boldsymbol{M}\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + \operatorname{Cov}(\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j)\right)\\
&\implies \widehat{\boldsymbol{\ell}} = \operatorname{Cov}(\rho_j, \boldsymbol{M}_j)\widehat{\boldsymbol{\ell}} + \operatorname{Cov}(\rho_j \boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j) + \rho \boldsymbol{M}\widehat{\boldsymbol{\ell}} + \operatorname{Cov}(\rho_j, \boldsymbol{M}^r_j)\widehat{\boldsymbol{r}} + \rho \boldsymbol{M}^r\widehat{\boldsymbol{r}}\\
&\qquad + (1-\rho)\left(\boldsymbol{M}\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + \operatorname{Cov}(\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j)\right)\\
&\implies \widehat{\boldsymbol{\ell}} = \operatorname{Cov}(\rho_j, \boldsymbol{M}_j)\widehat{\boldsymbol{\ell}} + \operatorname{Cov}(\rho_j \boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j) + \operatorname{Cov}(\rho_j, \boldsymbol{M}^r_j)\widehat{\boldsymbol{r}}\\
&\qquad + \boldsymbol{M}\widehat{\boldsymbol{\ell}} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + (1-\rho)\operatorname{Cov}(\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j)
\end{aligned}
$$

Rearranging yields the desired expression.

As for consumption, notice that:

$$
\begin{aligned}
\widehat{\boldsymbol{\ell}}^{NT}_j = \widehat{\boldsymbol{c}}^{NT}_j = \widehat{\boldsymbol{c}}_j &= \boldsymbol{M}_j\widehat{\boldsymbol{\ell}}_j + \boldsymbol{M}^r_j\widehat{\boldsymbol{r}}\\
&= \rho_j \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}^{NT}_j + (1-\rho_j)\boldsymbol{M}_j\widehat{\boldsymbol{\ell}}^{T}_j + \boldsymbol{M}^r_j\widehat{\boldsymbol{r}}
\end{aligned}
$$

$\widehat{\boldsymbol{\ell}}^{T}_j$ is equalized across counties. Moreover, $\widehat{\boldsymbol{c}}^{NT}_j = \widehat{\boldsymbol{c}}^{T}_j$. Hence $\widehat{\boldsymbol{\ell}}^{T}_j = \widehat{\boldsymbol{\ell}}^{NT}$ and:

$$
\widehat{\boldsymbol{\ell}}^{NT} = \boldsymbol{M}\widehat{\boldsymbol{\ell}}^{NT} + \boldsymbol{M}^r\widehat{\boldsymbol{r}} + \operatorname{Cov}(\rho_j \boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}^{NT}_j)
$$

From which it follows that:

$$
\operatorname{Cov}(\rho_j \boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}^{NT}_j) = \operatorname{Cov}(\rho_j \boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j(1-\rho_j)\boldsymbol{M}_j)\widehat{\boldsymbol{\ell}}^{NT} + \operatorname{Cov}(\rho_j \boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j\boldsymbol{M}^r_j)\widehat{\boldsymbol{r}}
$$

Combining the last two expressions gives the desired result.

B Extension to the New Keynesian Phillips Curve

B.1 Model features

New Keynesian Phillips Curves. In every county j, there are two sets of labor unions, one per sector. In every sector s, there is a continuum of unions indexed by $\iota \in [0,1]$. Labor services provided by different unions are packed into aggregate labor according to a CES aggregator with elasticity of substitution ϵ. Unions set their wage $w^s_{jt}(\iota)$ at any time t subject to quadratic utility costs to wage adjustment, governed by the parameter ψ. Thus, unions solve the following problem:

$$\max_{\{w^s_{jt+h}(\iota), \ell^s_{jt+h}(\iota)\}_{h\geq 0}} \quad \mathbb{E}_t \sum_{h\geq 0} \beta^{t+h} \left[u(c_{jt+h}) - v(\ell_{jt+h}) - \frac{\psi}{2} \left(\frac{w^s_{jt+h}(\iota)}{w^s_{jt+h-1}(\iota)} - 1 \right)^2 \right]$$
$$s.t. \quad \ell^s_{jt}(\iota) = \left(\frac{w^s_{jt}(\iota)}{w^s_{jt}} \right)^{-\epsilon} \ell^s_{jt}$$

Where c_{jt} and ℓ_{jt} are respectively aggregate consumption and aggregate labor supply in county j. Note that the union has preferences defined over the "average" or representative household of the county (this can of course be relaxed). The first-order condition to the union's problem reads:

$$u'(c_{jt}) \frac{\partial c_{jt}}{\partial w^s_{jt}(\iota)} - v'(\ell_{jt}) \frac{\partial \ell_{jt}}{\partial w^s_{jt}(\iota)} - \psi \left(\frac{w^s_{jt}(\iota)}{w^s_{jt-1}(\iota)} - 1 \right) \frac{1}{w^s_{jt-1}(\iota)} + \beta\psi \left(\frac{w^s_{jt+1}(\iota)}{w^s_{jt}(\iota)} - 1 \right) \frac{w^s_{jt+1}(\iota)}{w^s_{jt}(\iota)} \frac{1}{w^s_{jt}(\iota)} = 0$$

Which can be rewritten as:

$$u'(c_{jt}) w^s_{jt}(\iota) \frac{\partial c_{jt}}{\partial w^s_{jt}(\iota)} - v'(\ell_{jt}) w^s_{jt}(\iota) \frac{\partial \ell_{jt}}{\partial w^s_{jt}(\iota)} - \psi \pi^s_{jt}(\iota) \left(1 + \pi^s_{jt}(\iota) \right) + \beta\psi \pi^s_{jt+1}(\iota) \left(1 + \pi^s_{jt+1}(\iota) \right) = 0$$

Now notice:

$$\frac{\partial \ell_{jt}}{\partial w^s_{jt}(\iota)} = \frac{\partial \ell_{jt}}{\partial \ell^s_{jt}} \frac{\partial \ell^s_{jt}}{\partial \ell^s_{jt}(\iota)} \frac{\partial \ell^s_{jt}(\iota)}{\partial w^s_{jt}(\iota)}$$

From the households' perspective $\ell^s_{jt} = \int_0^1 \ell^s_{jt}(\iota)d\iota$ so $\frac{\partial \ell^s_{jt}}{\partial \ell^s_{jt}(\iota)} = 1$. Next, from $\ell^s_{jt}(\iota) = \left(\frac{w^s_{jt}(\iota)}{w^s_{jt}}\right)^{-\varepsilon} \ell^s_{jt}$, we have:

$$\frac{\partial \ell^s_{jt}(\iota)}{\partial w^s_{jt}(\iota)} = -\epsilon \frac{\ell^s_{jt}(\iota)}{w^s_{jt}(\iota)}$$

and:

$$\frac{\partial \ell_{jt}}{\partial \ell^s_{jt}} = \left(\frac{\ell^s_{jt}}{\alpha^s_j \ell_{jt}}\right)^{\frac{1}{\eta}} = \frac{w^s_{jt}}{w_{jt}}$$

Combining all of these together we have that:

$$\frac{\partial \ell_{jt}}{\partial w^s_{jt}(\iota)} = -\epsilon \frac{\ell^s_{jt}(\iota)}{w^s_{jt}(\iota)} \frac{w^s_{jt}}{w_{jt}}$$

As for the term $\frac{\partial c_{jt}}{\partial w^s_{jt}(\iota)}$, we can apply the envelope theorem and evaluate it as if all extra income was spent: $\frac{\partial c_{jt}}{\partial w^s_{jt}(\iota)} = \frac{\partial Z_{jt}}{\partial w^s_{jt}(\iota)}$. Then:

$$\frac{\partial c_{jt}}{\partial w^s_{jt}(\iota)} = \frac{1}{P_{jt}} \left(\frac{\partial w_{jt}}{\partial w^s_{jt}} \frac{\partial w^s_{jt}}{\partial w^s_{jt}(\iota)} \ell_{jt} + \frac{\partial \ell_{jt}}{\partial \ell^s_{jt}} \frac{\partial \ell^s_{jt}}{\partial \ell^s_{jt}(\iota)} \frac{\partial \ell^s_{jt}(\iota)}{\partial w^s_{jt}(\iota)} w_{jt} \right)$$

Note that:

$$\frac{\partial w_{jt}}{\partial w^s_{jt}} = \alpha^s_j \left(\frac{w^s_{jt}}{w_{jt}}\right)^{\eta} = \frac{\ell^s_{jt}}{\ell_{jt}}$$

And:

$$\frac{\partial w^s_{jt}}{\partial w^s_{jt}(\iota)} = \left(\frac{w^s_j(\iota)}{w^s_{jt}}\right)^{-\epsilon} = \frac{\ell^s_{jt}(\iota)}{\ell^s_{jt}}$$

Thus:

$$\frac{\partial c_{jt}}{\partial w_{jt}} = \frac{1}{P_{jt}} \left(\frac{\ell^s_{jt}}{\ell_{jt}} \frac{\ell^s_{jt}(\iota)}{\ell^s_{jt}} \ell_{jt} - \epsilon \frac{\ell^s_{jt}(\iota)}{w^s_{jt}(\iota)} \frac{w^s_{jt}}{w_{jt}} w_{jt} \right)$$

So now we can combine all of the above together and plug it into the FOC:

$$\pi^s_{jt}(\iota) + \pi^s_{jt}(\iota)^2 = \beta\left(\pi^s_{jt+1}(\iota) + \pi^s_{jt+1}(\iota)^2\right)$$
$$+ \frac{1}{\psi}\left[\frac{u'(c_{jt})}{P_{jt}}\left(w^s_{jt}(\iota)\ell^s_{jt}(\iota) - \epsilon w^s_{jt}\ell^s_{jt}(\iota)\right) + \epsilon v'(\ell_{jt})\ell^s_{jt}(\iota)\frac{w^s_{jt}}{w_{jt}}\right]$$

Imposing symmetry:

$$\pi^s_{jt} + {\pi^s_{jt}}^2 = \beta\left(\pi^s_{jt+1} + {\pi^s_{jt+1}}^2\right) + \frac{\epsilon}{\psi}\ell^s_{jt}\left[v'(\ell_{jt})\frac{w^s_{jt}}{w_{jt}} - \mu u'(c_{jt})\frac{w^s_{jt}}{P_{jt}}\right]$$

Where $\mu = \frac{\epsilon-1}{\epsilon}$. Evaluating the equation above at the 0 inflation steady-state:

$$v'(\ell_j)\ell_j = \mu u'(c_j)\frac{w_j}{P_j}\ell_j \tag{B.1}$$

Which is not s-dependent, meaning that the different unions' choices are consistent in the steady state. We then take a first-order approximation around the zero-inflation steady state:

$$\pi^s_{jt} = \beta\pi^s_{jt+1} + \frac{\epsilon}{\psi}\ell_j v'(\ell_j)\rho^s_j\left[\varphi\hat{l}_{jt} + \sigma\hat{c}_{jt} - \hat{w}^r_{jt}\right]$$

Where hat variables represent log-deviations from the steady state, $\rho^s_j = \frac{w^s_j\ell^s_j}{w_j\ell_j}$, $\varphi = \frac{v''(\ell_j)}{v'(\ell_j)}\ell_j$, $\sigma = -\frac{u''(c_j)}{u'(c_j)}c_j$ and w^r_{jt} is the real wage. We can iterate forward the equation above and write it in the sequence-space form as follows:

$$\boldsymbol{\pi}^s_j = \kappa^s_j\boldsymbol{K}\left[\varphi\widehat{\boldsymbol{\ell}}_j + \sigma\widehat{\boldsymbol{c}}_j - \widehat{\boldsymbol{w}}^r_j\right] \tag{B.2}$$

where:

$$\boldsymbol{K} \equiv \begin{bmatrix} 1 & \beta & \beta^2 & \cdots \\ 0 & 1 & \beta & \cdots \\ 0 & 0 & 1 & \cdots \\ \vdots & \vdots & \vdots & \ddots \end{bmatrix}$$

And $\kappa^s_j \equiv \frac{\epsilon}{\psi}\ell_j v'(\ell_j)\rho^s$. Finally, aggregate wage inflation in county j is given by:

$$\begin{aligned}\boldsymbol{\pi}^w_j =& \rho_j\boldsymbol{\pi}^{NT}_j + (1-\rho)\boldsymbol{\pi}^T_j \\ =& \kappa^w_j\boldsymbol{K}\left[\varphi\widehat{\boldsymbol{\ell}}_j + \sigma\widehat{\boldsymbol{c}}_j - \widehat{\boldsymbol{w}}^r_j\right]\end{aligned}$$

where:

$$\kappa_j^w \equiv \frac{\epsilon}{\psi}\ell_j v'(\ell_j)\left(\rho_j^2 + (1-\rho_j)^2\right)$$

Price inflation in county j, instead, reads as follows:

$$\begin{aligned}\boldsymbol{\pi}_j =& \xi_j \boldsymbol{\pi}_j^{NT} + (1-\xi_j)\boldsymbol{\pi}^T \\ =& \kappa_j \boldsymbol{K}\left[\varphi\widehat{\boldsymbol{\ell}}_j + \sigma\widehat{\boldsymbol{c}}_j - \widehat{\boldsymbol{w}}_j^r\right]\end{aligned}$$

Where $\xi_j \equiv \frac{P_j^{NT} c_j^{NT}}{P_j c_j}$ is the steady-state share of non-tradable consumption in county j and $\boldsymbol{\pi}^T \equiv \int_0^1 \boldsymbol{\pi}_j^T dj$ is the union-wide inflation rate for tradable goods and $\kappa_j \equiv \frac{\epsilon}{\psi}\ell_j v'(\ell_j)\left(\rho_j\xi_j + (1-\rho_j)(1-\xi_j)\right)$.

Monetary policy. When solving for the sticky-wage equilibrium, we assume monetary policy follows a standard Taylor rule given by:

$$i_t = r + \phi\pi_t + \varepsilon_t^i$$

where r is the steady-state real interest rate.

Equilibrium. We then modify our notion of equilibrium as follows:

Definition 6 (National Equilibrium with Sticky Wages)**.** *Given initial regional distributions $\{G_{j0}(b,e)\}_j$ over bonds b and idiosyncratic labor productivity e, and given exogenous paths of monetary shocks $\{\varepsilon_t^i\}_t$, an equilibrium is defined as a set of sequences $\{c_{jt}, c_{jt}^{NT}, c_{jt}^T, L_{jt}, L_{jt}^{NT}, L_{jt}^T, P_{jt}, P_{jt}^{NT}, P_{jt}^T, \pi_{jt}, \pi_{jt}^{NT}, \pi_{jt}^T, r_{jt}\}_{jt}$, union-wide nominal interest rates $\{i_t\}_t$, individual allocation rules $\{c_{jt}(b,e), b_{jt+1}(b,e)\}_{jt}$, and joint distributions over assets and productivity levels $\{G_{jt}(b,e)\}_{jt}$, such that households, unions, and firms in all counties optimize, the Taylor rule and the Fisher equation hold, and all markets clear:*

$$L_{jt}^{NT} = c_{jt}^{NT} \quad \textit{for all } j \tag{B.3}$$

$$L_{jt}^{T} = \int_0^1 c_{j't}^{T}(j)dj' \quad \textit{for all } j \tag{B.4}$$

$$\int_0^1 P_{jt}B_{jt}dj = \int_0^1 P_{jt}\int_0^1 b_{ijt}di\, dj \tag{B.5}$$

B.2 Regional Keynesian Cross with a New Keynesian Phillips Curve

Under sticky wages, the regional Keynesian cross cannot be expressed in terms of exogenous sequences only as the local price and wage responses will affect household's real income. Nonetheless, we can still derive an analogous expression accounting for the variation in prices and wages. In particular, the regional Keynesian cross for employment, i.e., the equivalent of (20) in the main text, becomes:

$$\begin{aligned}\widehat{\boldsymbol{\ell}}_j =& \rho_j \boldsymbol{M}^r_j \widehat{\boldsymbol{r}}_j + \rho_j \boldsymbol{M}_j \widehat{\boldsymbol{\ell}}_j + (1-\rho_j)\widehat{\boldsymbol{c}}^T + \rho_j \boldsymbol{M}^{\text{cap}}_j \boldsymbol{\pi}^{\text{surprise}}_j \\ &+ \rho_j \boldsymbol{M}_j \left[(\rho_j - \xi_j)\left(\widehat{\boldsymbol{w}}^{NT}_j - \widehat{\boldsymbol{p}}^T\right) + (1-\rho_j)\left(\widehat{\boldsymbol{w}}^T_j - \widehat{\boldsymbol{p}}^T\right)\right] \\ &- \nu\rho_j(1-\xi_j)\left(\widehat{\boldsymbol{w}}^{NT}_j - \widehat{\boldsymbol{p}}^T\right) - \theta(1-\rho_j)\left(\widehat{\boldsymbol{w}}^T_j - \widehat{\boldsymbol{p}}^T\right)\end{aligned} \tag{B.6}$$

And the corresponding one for consumption becomes:

$$\begin{aligned}\widehat{\boldsymbol{c}}_j =& \boldsymbol{M}^r_j \widehat{\boldsymbol{r}}_j + \rho_j \boldsymbol{M}_j \widehat{\boldsymbol{c}}_j + (1-\rho_j)\boldsymbol{M}_j \widehat{\boldsymbol{c}}^T + \boldsymbol{M}^{\text{cap}}_j \boldsymbol{\pi}^{\text{surprise}}_j \\ &+ \boldsymbol{M}_j \left(((\rho_j - \xi_j) - \nu\rho_j(1-\xi_j))\left(\widehat{\boldsymbol{w}}^{NT}_j - \widehat{\boldsymbol{p}}^T\right) + (1-\rho_j)(1-\theta)\left(\widehat{\boldsymbol{w}}^T_j - \widehat{\boldsymbol{p}}^T\right)\right)\end{aligned} \tag{B.7}$$

where $\widehat{\boldsymbol{p}}^T - \widehat{\boldsymbol{w}}^T_j$ denotes the relative price of imports over exports, i.e., the terms of trade, $\boldsymbol{M}^{\text{cap}}$ is the consumption Jacobian to surprise capital gains induced by unanticipated inflation and $\boldsymbol{\pi}^{\text{surprise}}_j$ represents surprise inflation.[44] Compared to the rigid wages version in the main text, this equation features a real income channel akin to the one in Auclert et al. (2021b) coming from the relative change of local wages and local prices. Moreover, it also features two expenditure switching margins: the first arises from the relative price change between non-tradable goods and the tradable bundle, while the second occurs within tradable goods, driven by the change in relative prices between local and national tradable goods. Together with the New Keynesian Phillips Curve, the Taylor rule and the Fisher equation, (B.7) pin down the employment response to monetary shocks.

Integrating (B.7), we can obtain a characterization of the National Keynesian Cross for employment under sticky wages, i.e., the equivalent of (27) in the main text:

$$\begin{aligned}\widehat{\boldsymbol{\ell}} =& \Big(\boldsymbol{M} + \text{Cov}(\rho_j, \boldsymbol{M}_j)\Big)\widehat{\boldsymbol{\ell}} + \Big(\boldsymbol{M}^r + \text{Cov}(\rho_j, \boldsymbol{M}^r_j)\Big)\widehat{\boldsymbol{r}} + \Big(\boldsymbol{M}^{\text{cap}} + \text{Cov}(\rho_j, \boldsymbol{M}^{\text{cap}}_j)\Big)\widehat{\boldsymbol{\pi}}^{\text{surprise}} \\ &+ \text{Cov}((1+\rho_j-\rho))\boldsymbol{M}_j, \widehat{\boldsymbol{\ell}}_j) + \text{Cov}((1+\rho_j-\rho))\boldsymbol{M}^r_j, \widehat{\boldsymbol{r}}_j) + \text{Cov}((1+\rho_j-\rho))\boldsymbol{M}^{\text{cap}}_j, \widehat{\boldsymbol{\pi}}^{\text{surprise}}_j)\end{aligned}$$

[44] Because we focus on perfect-foresight MIT shocks, $\boldsymbol{\pi}^{\text{surprise}}_j$ is a vector of zeros, except for the first entry which contains on-impact inflation, π_{0j}.

$$+ (\widehat{\boldsymbol{w}} - \widehat{\boldsymbol{p}}) (-\nu + \boldsymbol{M}) + \mathrm{Cov}(\widehat{\boldsymbol{w}}_j - \widehat{\boldsymbol{p}}_j, \boldsymbol{M}_j(1 + \rho_j - \rho)) + (\nu - \theta) \left(\widehat{\boldsymbol{w}}^T - \widehat{\boldsymbol{p}}^T - \mathrm{Cov}(\widehat{\boldsymbol{w}}_j^T, \rho_j) \right) \tag{B.8}$$

Note that (B.8) nests (27) in the main text in the limit of rigid wages. Moreover, under the case $\nu = \theta$—which nests the Cole-Obstfeld calibration—the last term in (B.8) vanishes.

Similarly, we can obtain the following characterization of the National Keynesian Cross for consumption under sticky wages, i.e., the equivalent of (26) in the main text:

$$\begin{aligned}
\widehat{\boldsymbol{c}} = &\left(\boldsymbol{M} + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, (1-\rho_j) \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j) \right) \widehat{\boldsymbol{c}} + \left(\boldsymbol{M} + \mathrm{Cov}(\boldsymbol{M}_j, \frac{\widehat{\boldsymbol{w}}_j^r}{\widehat{\boldsymbol{w}}^r}) + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j \frac{\widehat{\boldsymbol{w}}_j^r}{\widehat{\boldsymbol{w}}^r}) \right) \widehat{\boldsymbol{w}}^r \\
&+ \left(\boldsymbol{M}^r + \mathrm{Cov}(\boldsymbol{M}_j^r, \frac{\widehat{\boldsymbol{r}}_j}{\widehat{\boldsymbol{r}}}) + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j^r \frac{\widehat{\boldsymbol{r}}_j}{\widehat{\boldsymbol{r}}}) \right) \widehat{\boldsymbol{r}} \\
&+ \left(\boldsymbol{M}^{\text{cap}} + \mathrm{Cov}(\boldsymbol{M}_j^{\text{cap}} \frac{\widehat{\boldsymbol{\pi}}_j^{\text{surprise}}}{\widehat{\boldsymbol{\pi}}^{\text{surprise}}}) + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j^{\text{cap}} \frac{\widehat{\boldsymbol{\pi}}_j^{\text{surprise}}}{\widehat{\boldsymbol{\pi}}^{\text{surprise}}}) \right) \widehat{\boldsymbol{\pi}}^{\text{surprise}} \\
&- \theta \left[\boldsymbol{M} \left(1 - \rho - \mathrm{Cov}(\rho_j, \boldsymbol{M}_j)\right) \left(\widehat{\boldsymbol{w}}^T - \widehat{\boldsymbol{p}}^T \right) + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, (1-\rho_j) \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j (\widehat{\boldsymbol{w}}_j^T - \widehat{\boldsymbol{p}}^T)) \right] \\
&- \nu \Big[(\boldsymbol{M}(\rho - \xi) + \mathrm{Cov}(\rho_j, \boldsymbol{M}_j))(\widehat{\boldsymbol{w}}^{NT} - \widehat{\boldsymbol{p}}^T) + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, (1-\xi_j)(\widehat{\boldsymbol{w}}_j^{NT} - \widehat{\boldsymbol{p}}^T)) \\
&- \boldsymbol{M} \, \mathrm{Cov}(\xi_j, (\widehat{\boldsymbol{w}}_j^{NT} - \widehat{\boldsymbol{p}}^T)) + \mathrm{Cov}(\rho_j \boldsymbol{M}_j, \rho_j (1-\xi_j) \boldsymbol{\mathcal{M}}_j \boldsymbol{M}_j (\widehat{\boldsymbol{w}}_j^{NT} - \widehat{\boldsymbol{p}}^T)) \Big]
\end{aligned} \tag{B.9}$$

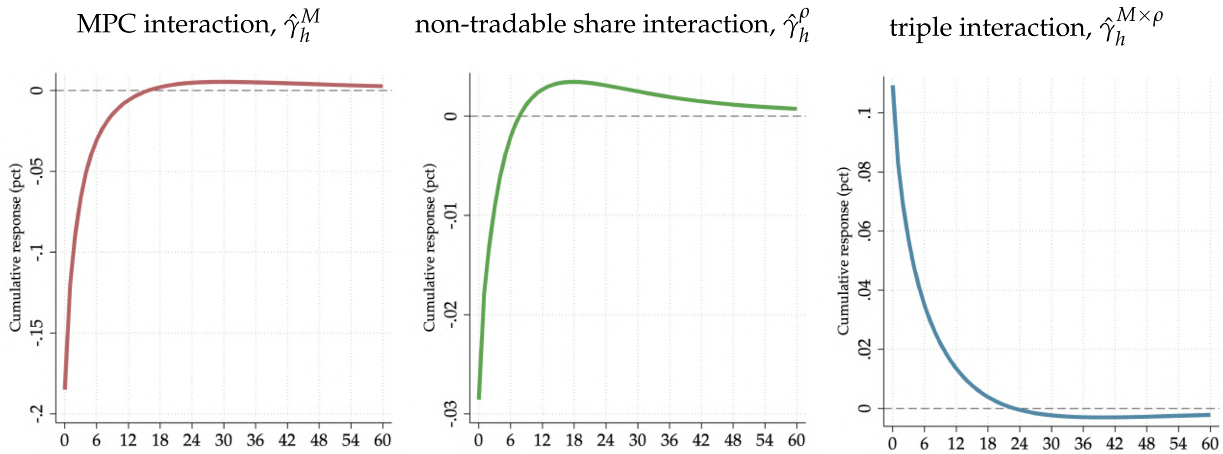
(B.9) also nests (26) in the main text in the limit case of rigid wages. Relative to (26), it features both a nationwide real wage channel as well as a term capturing the covariance between changes in real wages and the objects governing the regional Keynesian multiplier, as well as terms capturing the consumption response generated by surprise inflation and expenditure switching.

B.3 Quantitative results

In this section, we replicate our main quantitative results under the assumption of sticky—rather than rigid—wages. We maintain exactly the same calibration as in Section 4.2, with the exception that we now parameterize the parameter ψ, governing wage adjustment costs, to hit a slope of the New Keynesian Wage Phillips Curve κ of 0.006, as in Auclert et al. (2021b). We also set a Taylor coefficient ϕ equal to 1.25.

First, Figure B.1 shows model-based local projection estimates in a fashion analogous to Figure 7 in the main text. As before, we consider a monetary shock normalized to increase national employment by 1% on impact. Second, Figure B.2 presents aggregate

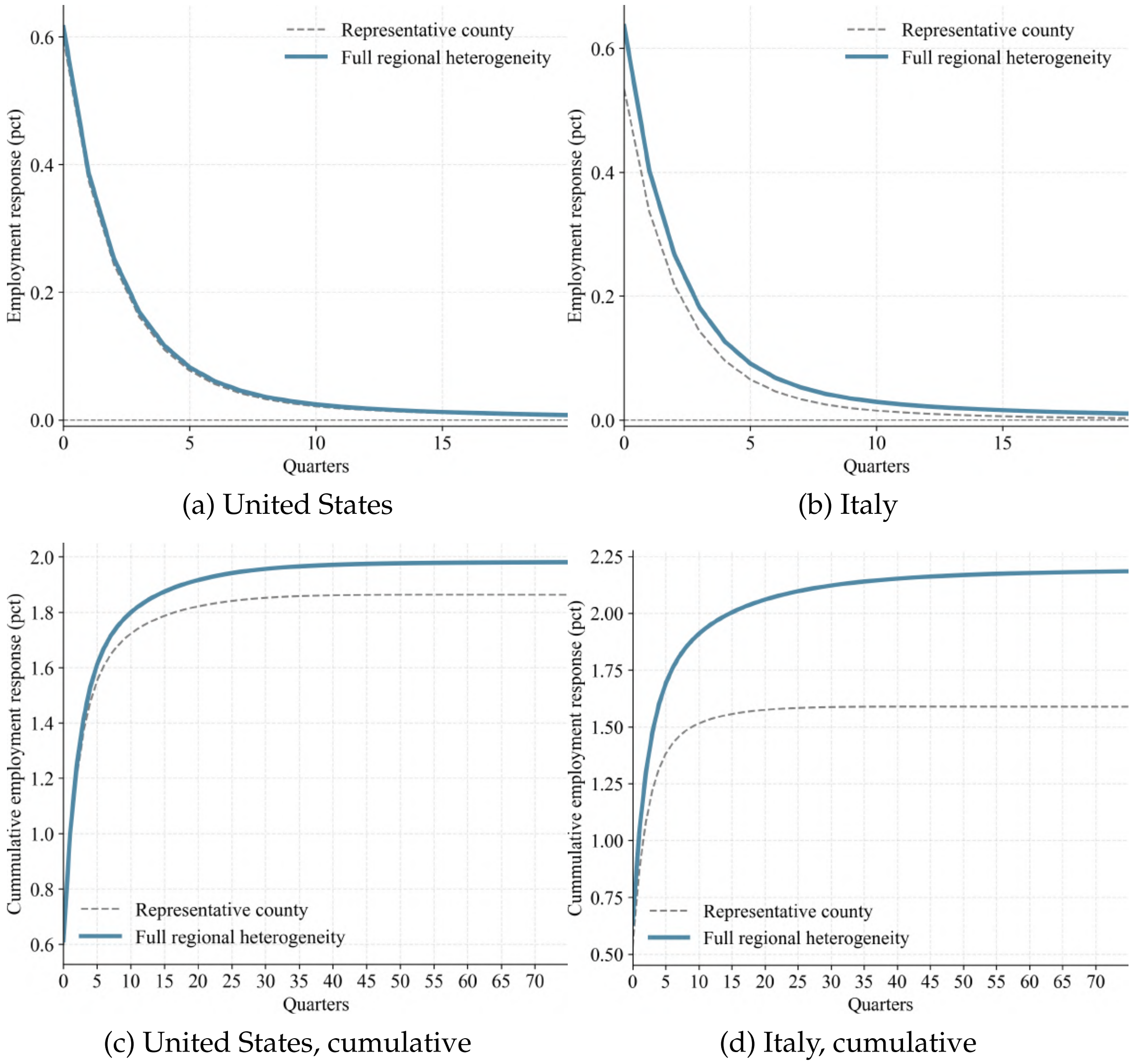
Figure B.1: Model-based local projections with sticky wages



Note: model-based impulse response functions capturing the coefficients in (32) *in response to an expansionary monetary policy shock that raises average employment by* 1% *on impact with quarterly persistence of 0.61.*

employment responses to a monetary policy shock for both the American and Italian economies, for the baseline and special cases with full regional heterogeneity and a representative county, respectively, and in the form of standard and cumulative impulse responses. In this case, same as in the main text, the monetary shock is a 1 p.p. (annualized) decline in r_t with quarterly persistence of 0.61. As can be seen from both Figure B.1 and Figure B.2, our main results remain quantitatively unchanged.

Figure B.2: Amplification with the New Keynesian Phillips Curve

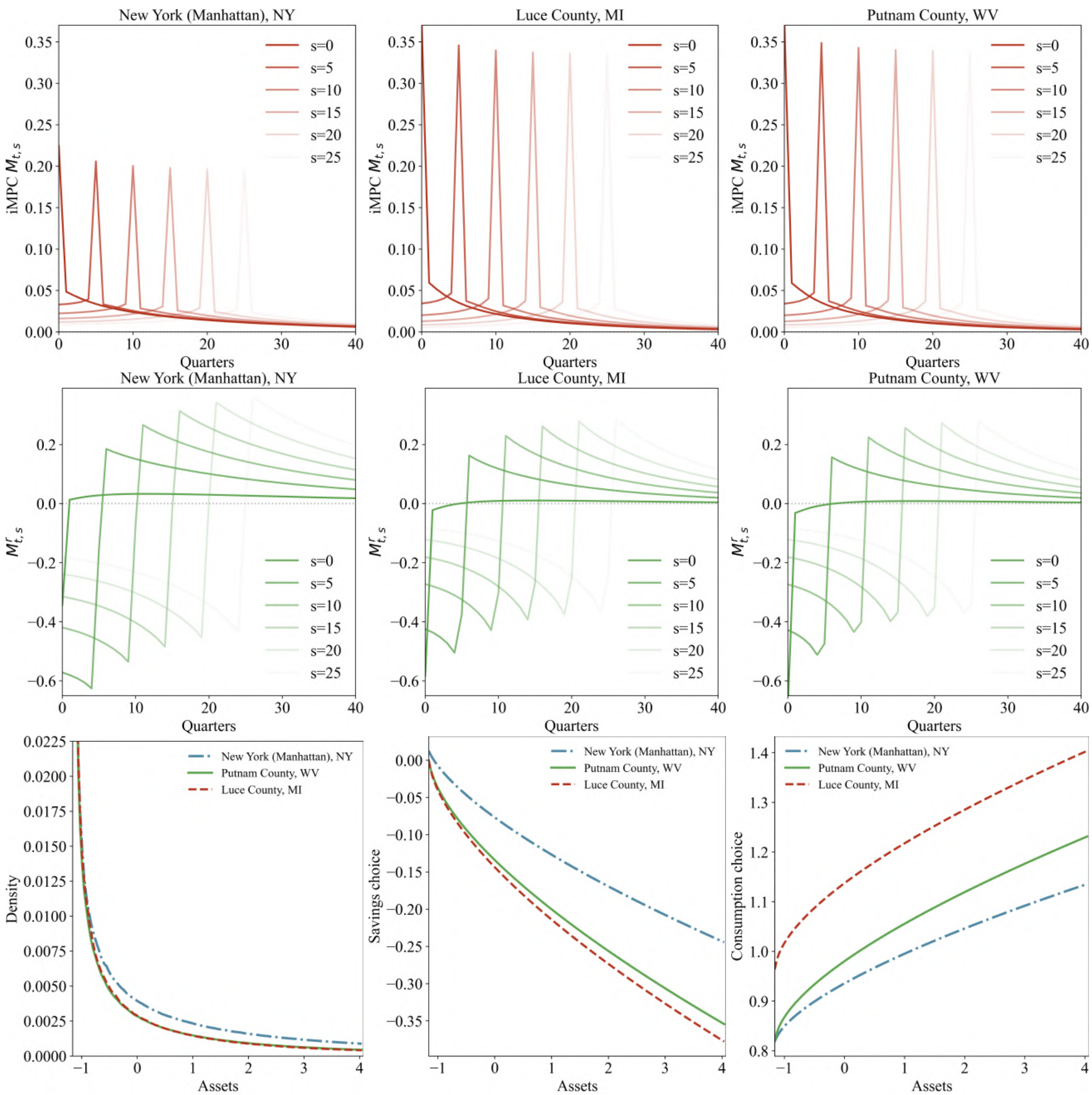


Note: employment responses to a monetary shock that decreases i_t by 1 p.p. (annualized) on impact with quarterly persistence of 0.61. Calibration of the steady state is identical to that used in the main text.

C Additional Figures and Tables

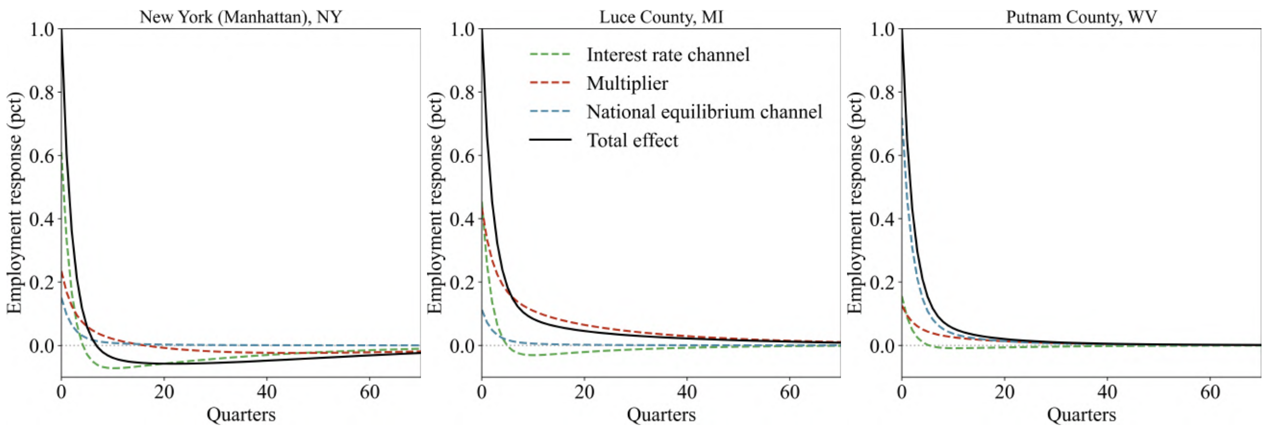
C.1 Additional model figures and tables

Figure C.1: Jacobians, distributions and policy functions



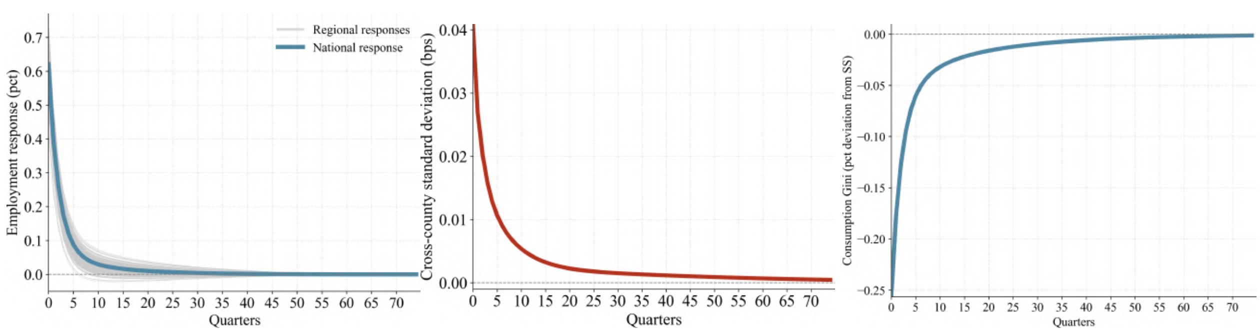
Note: The first three panels represent the columns of the iMPC matrix $\boldsymbol{M}_j[t,s]$ for different values of s for the three example counties. The middle three panels represent the columns of the matrix $\boldsymbol{M}^r_j[t,s]$ for different values of s for the three example counties. The last three panels represent the stationary asset distributions as well as savings and consumption policy functions in each of the three counties. Policy functions are displayed for the median income state.

Figure C.2: Decomposing the heterogeneous response to monetary policy, long horizon



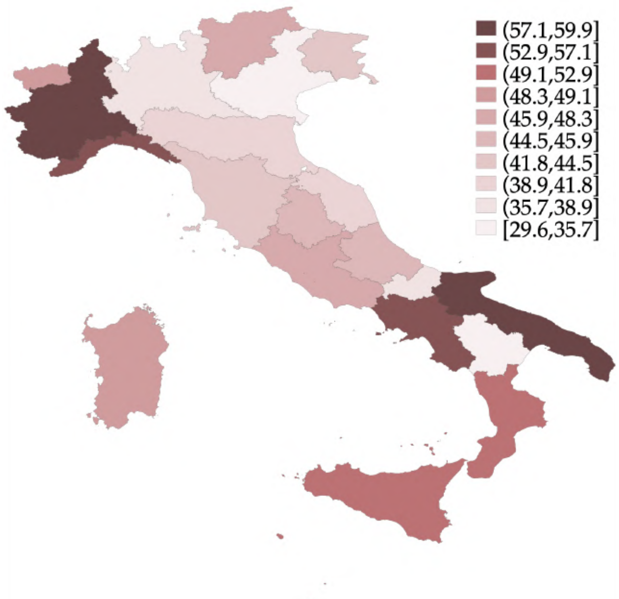
Note: decomposition of the employment response to a monetary shock that decreases r_t by 1 p.p. (annualized) on impact with quarterly persistence of 0.61.

Figure C.3: Regional heterogeneity and inequality, longer horizon

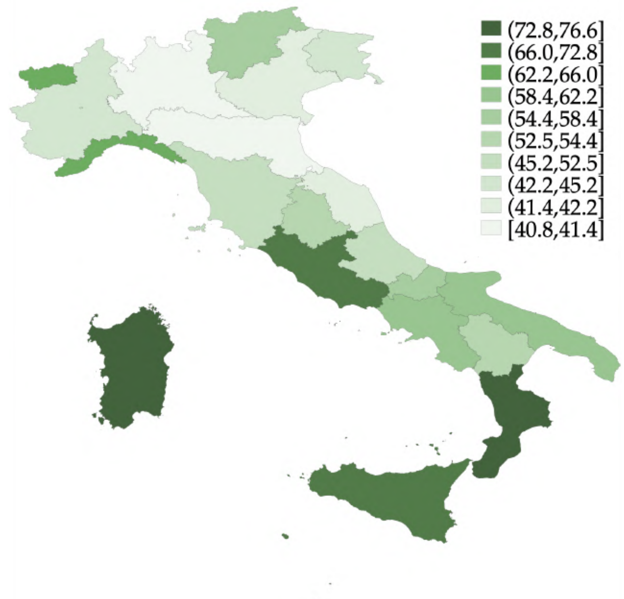


Note: employment responses to a monetary shock that decreases r_t by 1 p.p. (annualized) on impact with quarterly persistence of 0.61. The left panel shows all county-level responses and the national response. The central panel plots the cross-county standard deviation of percentage output deviations (effectively, the horizon-specific standard deviation of the grey respones in the left panel of the graph). The right panel of the figure shows the extension of Figure 11 to 75 quarters.

Figure C.4: The geography of Italian MPCs and trade openness



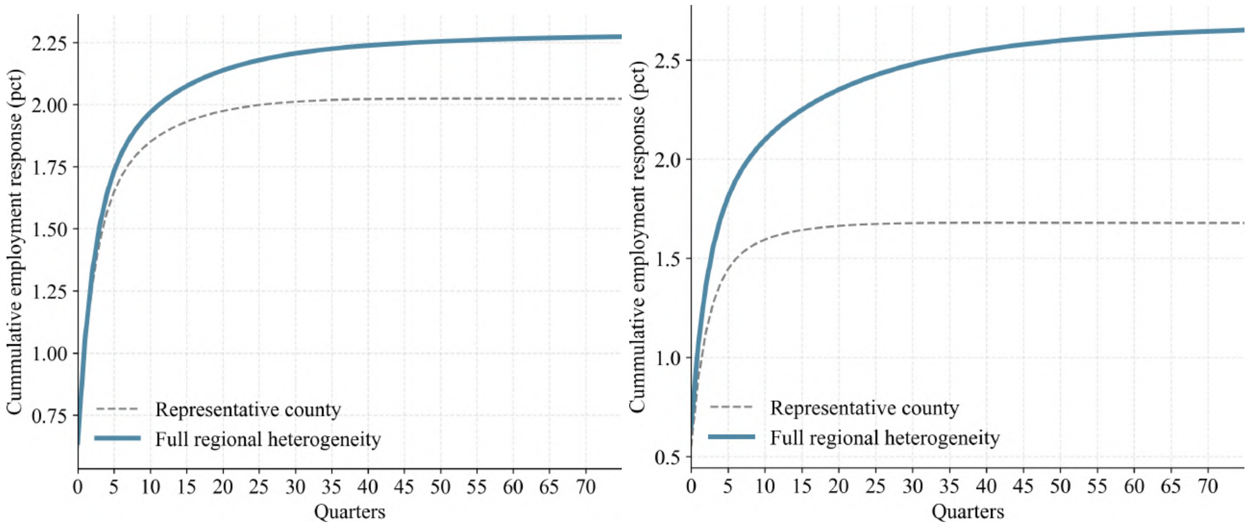
(a) MPC



(b) Non-tradable share

Note: the left and right panels plot the geographic distributions of MPCs and non-tradable shares in Italy, respectively.

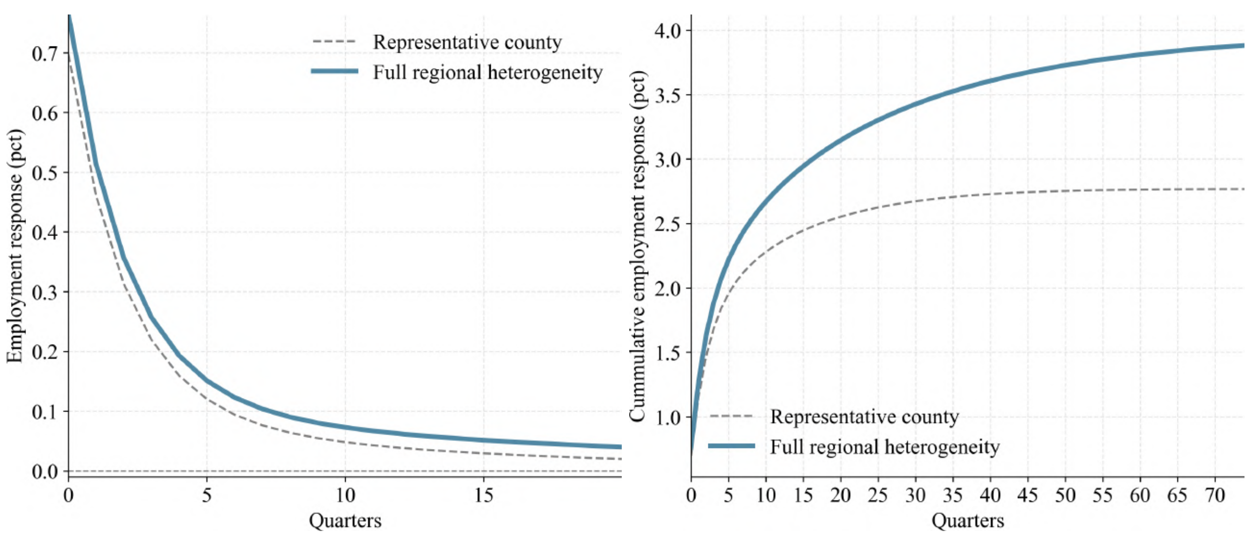
Figure C.5: Regional heterogeneity and aggregate amplification—cumulative response



(a) United States (b) Italy

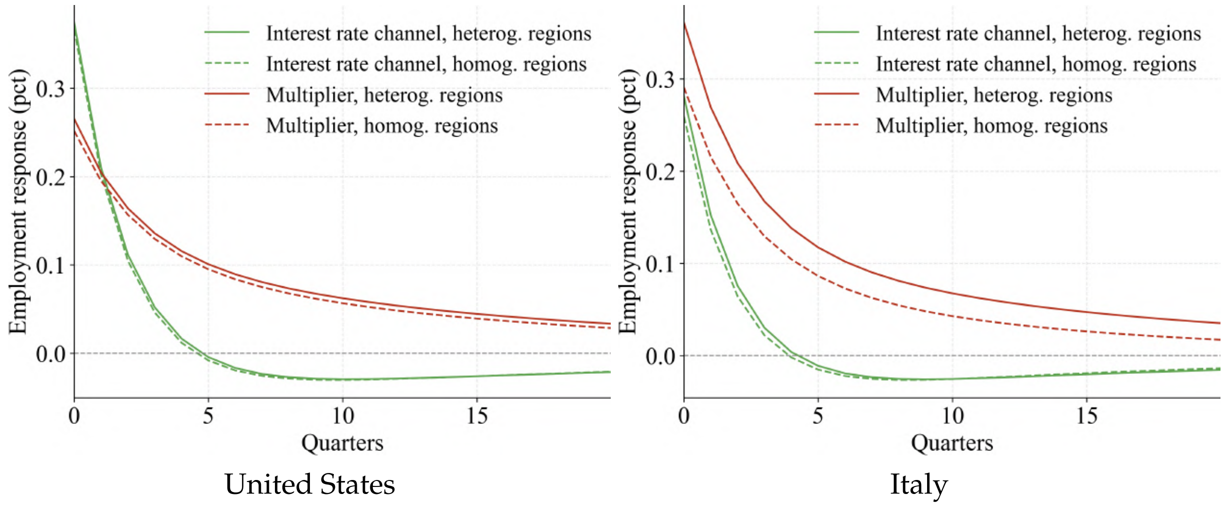
Note: the figure plots cumulative employment responses to a monetary shock that decreases r_t by 1 p.p. (annualized) on impact with quarterly persistence of 0.61, in the full model vs. the representative-county case. The representative county represents the response of the re-calibrated economy with no regional heterogeneity while the full regional heterogeneity response refers to the calibrated heterogeneous-county model from the main text.

Figure C.6: Regional heterogeneity and aggregate amplification—Italy and annual MPCs



Note: this figure plots the equivalent of Figure 12, panel (b), with the Italian economy now calibrated to yearly, rather than quarterly, regional MPCs.

Figure C.7: Decomposition of the National Keynesian cross



Note: this figure leverages the national Keynesian cross (Proposition 3) to decompose the sources of aggregate amplification into direct (interest rate channel) and indirect (multiplier) effects.

Table C.1: Model calibration of county-specific parameters

Statistic	**Model**	**Data**
Mean of MPC_j	0.308	0.302
Mean of ρ_j	0.604	0.608
Standard deviation of MPC_j	0.038	0.039
Standard deviation of ρ_j	0.150	0.157
Covariance between MPC_j and ρ_j	-0.002	-0.002

Note: select moments in the model and the U.S. data.

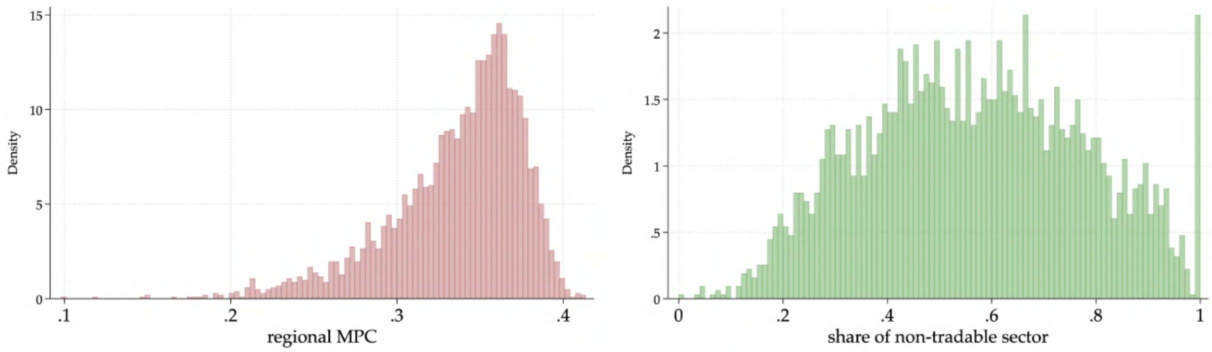
Table C.2: Parametrization of the representative counties

No.	**Parameters**		**Targets**		**No.**	**Parameters**		**Targets**	
	α_j	β_j	ρ_j	MPC_j		α_j	β_j	ρ_j	MPC_j
1.	0.67397	0.98468	0.67199	0.18569	11.	0.64338	0.97500	0.63831	0.48259
2.	0.38294	0.98283	0.39149	0.24659	12.	0.52222	0.97400	0.52331	0.50272
3.	0.43123	0.98085	0.43760	0.28723	13.	0.79598	0.97245	0.78331	0.53342
4.	0.41080	0.98055	0.41775	0.29297	14.	0.70015	0.97237	0.69170	0.53491
5.	0.40037	0.97988	0.40746	0.34063	15.	0.41724	0.97200	0.42263	0.54123
6.	0.59427	0.97980	0.59304	0.34218	16.	0.58501	0.97169	0.58264	0.54688
7.	0.51345	0.97826	0.51572	0.37512	17.	0.70324	0.97130	0.69451	0.55363
8.	0.46356	0.97635	0.46764	0.41420	18.	0.72920	0.96710	0.71884	0.66951
9.	0.39987	0.97635	0.40630	0.41414	19.	0.62062	0.96695	0.61599	0.67162
10.	0.50310	0.97600	0.50539	0.42088	20.	0.53012	0.96650	0.53028	0.67773

Note: this table shows the results from the first step of the model calibration procedure that yields the calibrated steady state of a national economy comprising 20 interlinked representative incomplete-market economies. For each panel, the first two columns summarize model parameters and the last two columns present the relevant data moments.

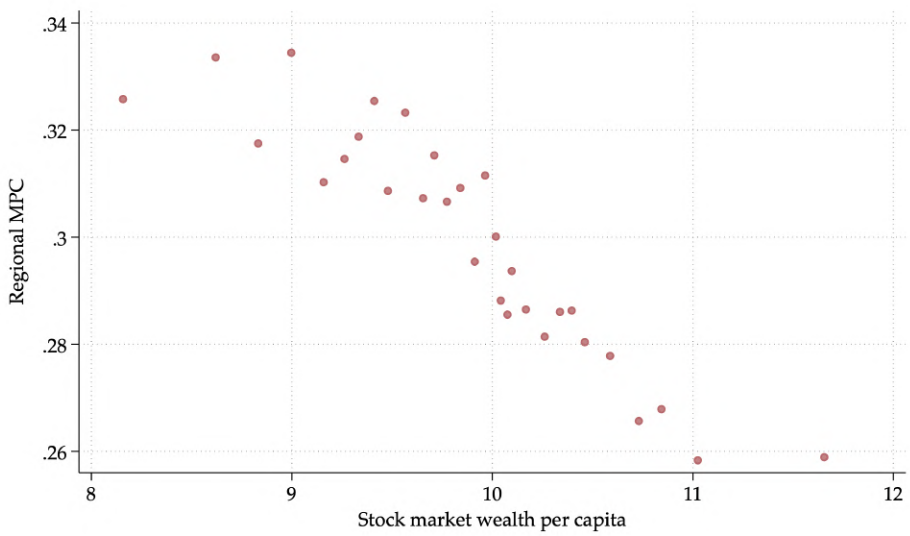
C.2 Additional empirical figures and tables

Figure C.8: Histograms of regional U.S. MPCs and non-tradable shares



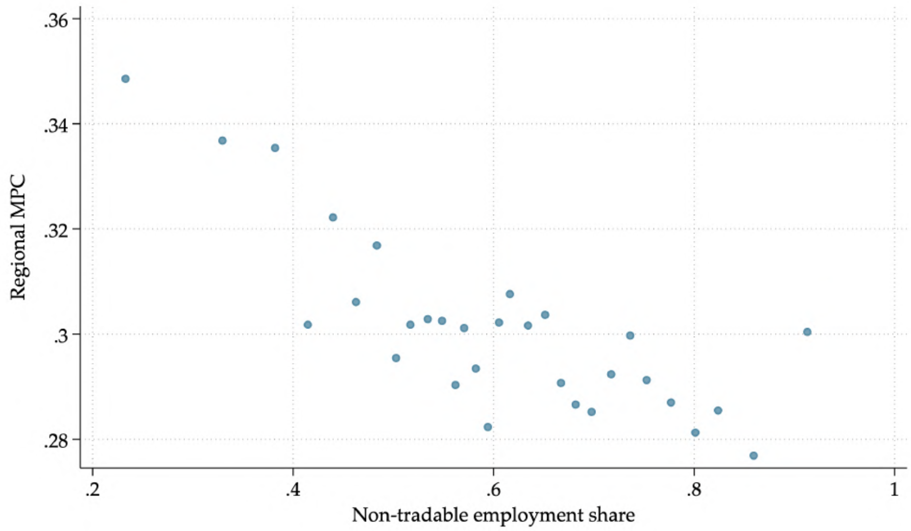
Note: histograms of U.S. county-level MPCs and non-tradable shares for the year 2016. See main text for details on the measurement methodology.

Figure C.9: Validating regional U.S. MPCs with stock market wealth



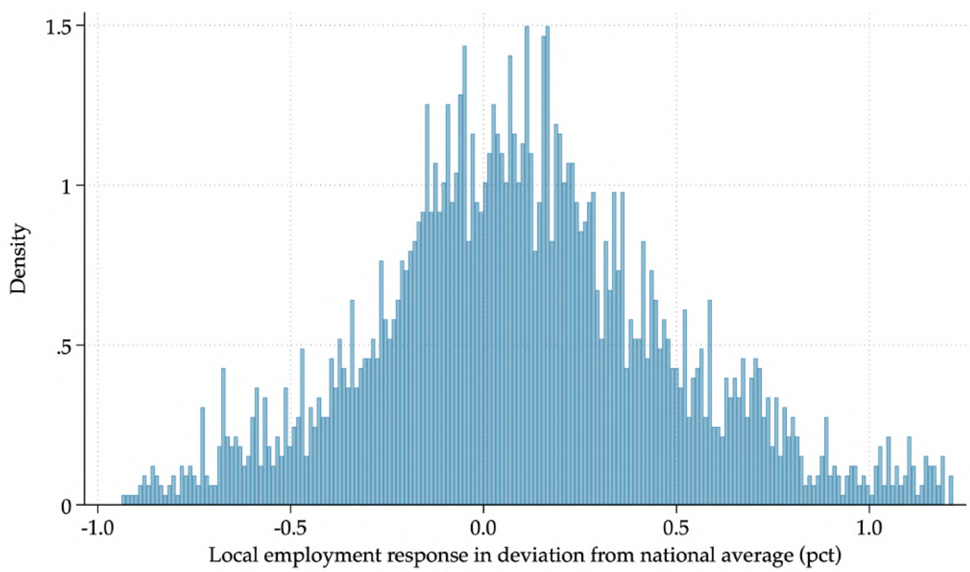
Note: population-weighted binned scatterplot of the U.S. county-level measure of MPCs against county-level stock market wealth per capita from Chodorow-Reich et al. (2021). Every bin covers approximately 3.3% of the U.S. population.

Figure C.10: The joint distribution of U.S. MPCs and non-tradable employment



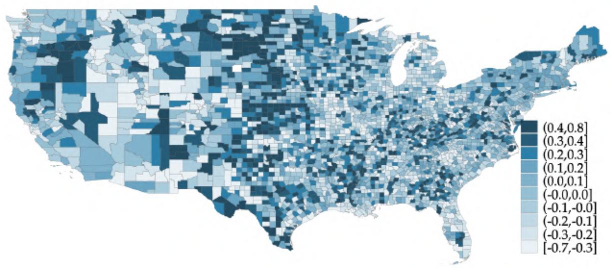
Note: population-weighted binned scatterplot of the county-level measure of MPCs against county-level non-tradable share of employment. Every bin covers approximately 3.3% of the U.S. population.

Figure C.11: Regional heterogeneity in the effects of U.S. monetary policy

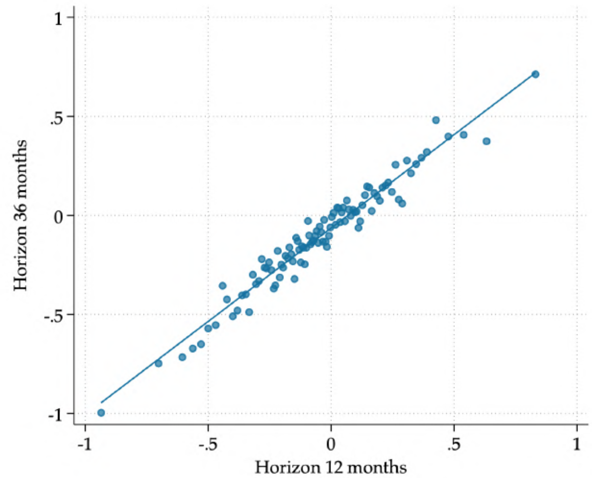


Note: This figure plots the histogram of the 3-year ahead county-specific cumulative employment responses to a 1 standard deviation expansionary monetary policy shock, $\hat{\beta}_{j,36}$, estimated from the panel local projection (31). The coefficients are in percentage points and represent deviations from the (population-weighted) average response.

Figure C.12: Regional responses to monetary shocks—comparing horizons



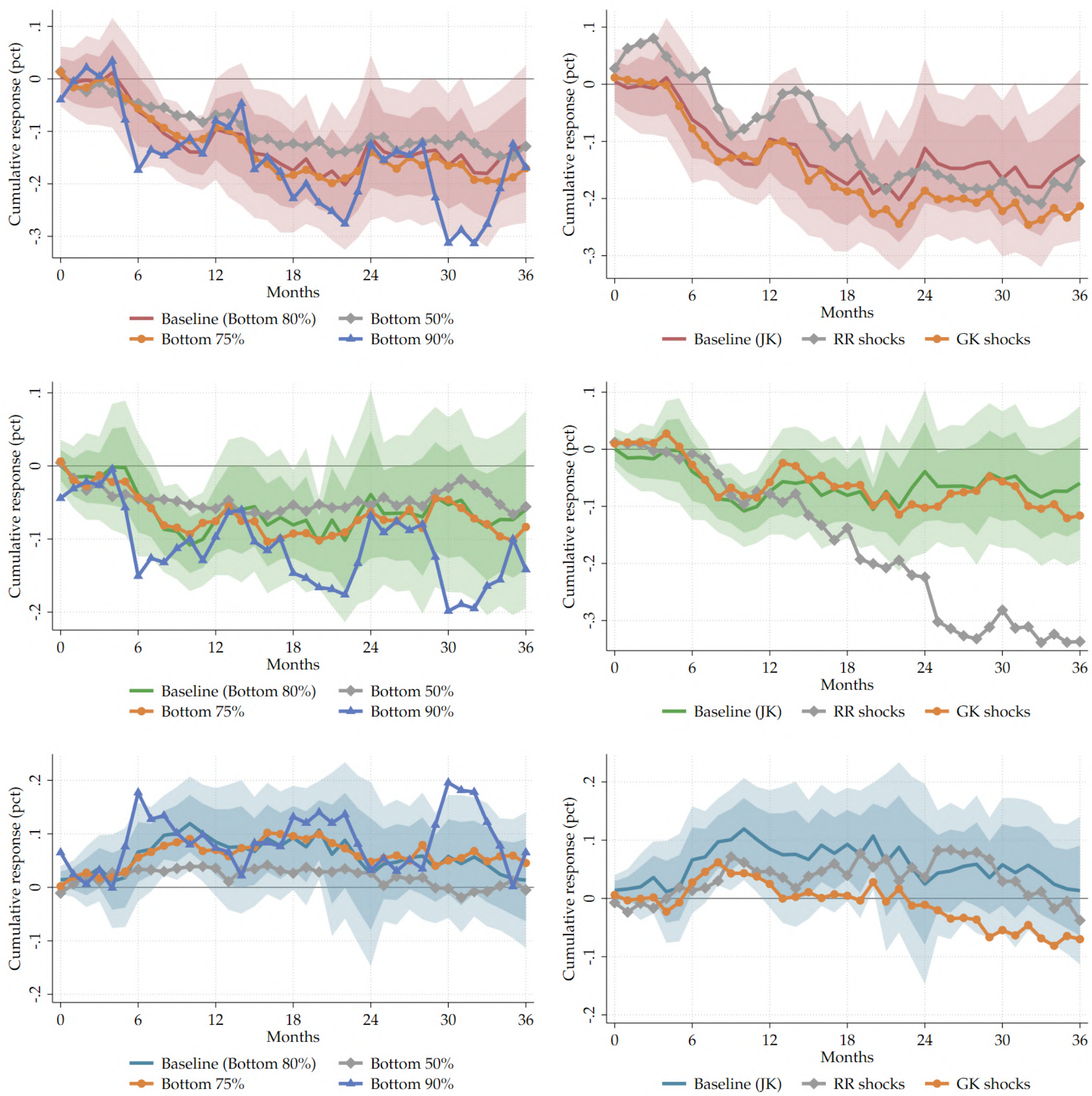
(a) 12-month ahead



(b) Horizon 12 vs horizon 36

Note: These figures plot results from the horizon robustness tests for the regression of n-month ahead county-specific cumulative employment responses to a 1 standard deviation expansionary monetary policy shock $\hat{\beta}_{j,n}$*, estimated from the panel local projection* (31)*. Panel (a) displays the geography of estimates 12 months ahead. Panel (b) presents the scatter plot of estimates from 12- and 36-month specifications with a line of best fit.*

Figure C.13: Robustness to different baseline groups and monetary surprises



Note: Robustness of empirical impulse responses for different baseline groups (the left column) and different monetary policy surprises (the right column). The top, middle, and bottom panels represent the MPC interactions (β_h^M), the non-tradable interactions (β_h^ρ), and the triple interactions ($\beta_h^{M\times\rho}$), respectively. A baseline group is defined as the part of the MPC (ρ) distribution such that D_j^M (D_j^ρ) is equal to zero. We conduct robustness for thresholds at the 50^{th}, 75^{th} and 90^{th} percentiles. As monetary surprises, the Romer and Romer (2000) and Gertler and Karadi (2015) shocks are used.

Table C.3: Estimates from regression (29)

Variable	Coeff	Variable	Coeff	Variable	Coeff	Variable	Coeff	Variable	Coeff
Black	-.192	HS diploma	.037	25-44 y.o.	.474	10-20k	.046	75-100k	.039
	(.05)		(.08)		(.13)		(.07)		(.07)
Asian	.005	College	-.024	45-64 y.o.	.405	20-30k	.114	100-150k	-.114
	(.08)		.08		.13		.07		.07
Hispanic	-.035	Bachelor	-.031	$>$ 65 y.o.	.413	30-40k	-.067	150-200k	-.166
	(.05)		(.08)		(.13)		(.07)		(.09)
Other	-.048	Masters	-.066	Unempl.	.076	40-60k	.095	$>$200k	-.229
	(.06)		(.09)		(.08)		(.08)		(.09)
Fin. inc. >0	-0.100	Rent house	.068	Other empl.	-.052	60-75k	.007	Constant	-.024
	(.04)		(.04)		(.04)		(.08)		(.157)
Observations	1170								
R-squared	.075								

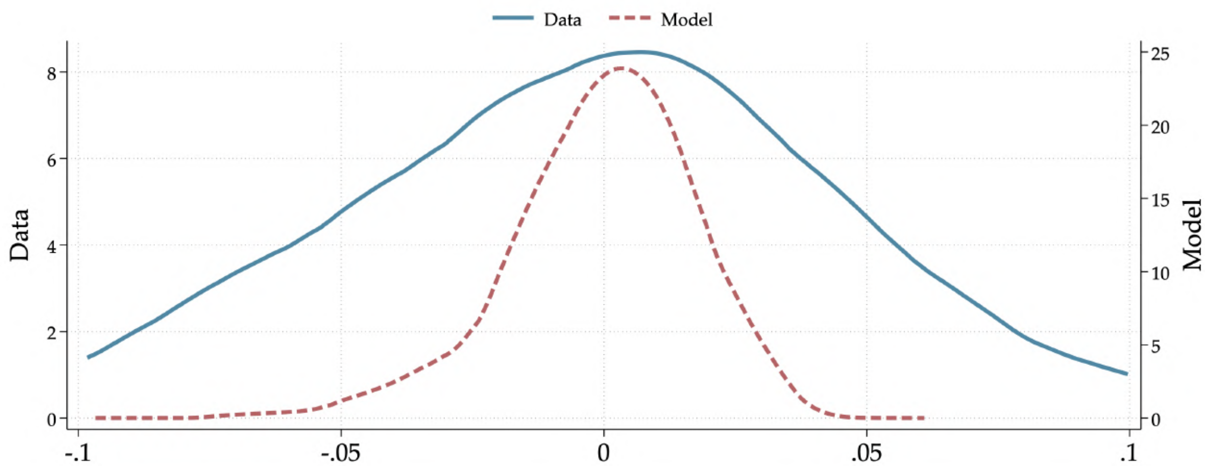
Note: This table reports the estimates from (29), which is the first step of the MPC computation procedure.

Table C.4: Comparison of distributional moments between U.S. and Italy

Statistic	US	Italy
Mean of MPC_j	0.302	0.451
Mean of ρ_j	0.608	0.55
Standard deviation of MPC_j	0.039	0.139
Standard deviation of ρ_j	0.157	0.117
Correlation between MPC_j and ρ_j	-0.313	0.451

Note: this table compares regional differences between the Italian and U.S. economies. It reports the first and second moments of regional distributions of MPCs and non-tradable shares, as well as the within-country correlations between MPCs and non-tradable shares across regions.

Figure D.1: Distributional effects of monetary policy shocks



Note: demeaned distribution of peak employment responses in the model and in the data, subject to the EB adjustment described in the text. The shock is scaled so that the average response is the same in the data and in the model.

D Empirical Bayes shrinkage

In order to estimate the "true" between-county heterogeneity in the response of employment to monetary shocks, β_j, we use an empirical-Bayes (EB) shrinkage procedure as described in Chetty et al. (2014) and more closely related Kleven (2024). For each county we have an estimate $\hat{\beta}_j$ and a corresponding standard error $\mathrm{SE}(\hat{\beta}_j)$. Our goal is to infer each group's latent β_j while also estimating the overall heterogeneity across groups.

We posit a two-stage hierarchical model:

$$\beta_j \sim N(0,\ \tau^2),$$

$$\hat{\beta}_j \,|\, \beta_j \ \sim \ N\Big(\beta_j,\ \mathrm{SE}(\hat{\beta}_j)^2\Big).$$

Hence, τ^2 represents the between-group variance in the true effects β_j, while $\mathrm{SE}(\hat{\beta}_j)^2$ captures the sampling variance of the observed estimate $\hat{\beta}_j$. Using an empirical-Bayes approach, we first estimate τ^2 from the observed $\hat{\beta}_j$. One method-of-moments estimator is:

$$\widehat{\tau}^2 \ = \ \max\left\{0,\ \frac{1}{N-1}\sum_{j=1}^{N}(\hat{\beta}_j - \overline{\hat{\beta}})^2 \ - \ \frac{1}{N}\sum_{j=1}^{N}\mathrm{SE}(\hat{\beta}_j)^2\right\},$$

where $\overline{\hat{\beta}} = \frac{1}{N}\sum_{j=1}^{N}\hat{\beta}_j$. Given $\widehat{\tau}^2$, we form the EB posterior mean for each group's true effect

β_j. If the prior is centered at 0 with variance $\widehat{\tau}^2$, the posterior mean (β_j^*) can be written as:

$$\beta_j^* \;=\; \frac{\widehat{\tau}^2}{\widehat{\tau}^2 + \mathrm{SE}(\hat{\beta}_j)^2}\,\hat{\beta}_j.$$

This "shrinkage factor" $\dfrac{\widehat{\tau}^2}{\widehat{\tau}^2 + \mathrm{SE}(\hat{\beta}_j)^2}$ determines how much $\hat{\beta}_j$ is shrunk toward 0. Larger sampling variances ($\mathrm{SE}(\hat{\beta}_j)$) yield more shrinkage. The final quantity $\widehat{\tau}^2$ (or $\widehat{\tau}^2$) measures the estimated "true" variance among the β_j values, after removing sampling noise. The posterior means β_j^* are then stabilized (or regularized) estimates of each group's true effect. The outcome of this procedure is shown in Figure D.1 that plots the full distribution of regional responses to monetary policy in the model and the data.