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Monetary policy and regional inequality

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Abstract

We study the impact of monetary policy on regional inequality using granular data on economic activity at the city- and county-level in Europe. We document pronounced heterogeneity in the regional patterns of monetary policy transmission. The output response to monetary policy shocks is stronger and more persistent in poorer regions, with the difference becoming particularly pronounced in the tails of the distribution. Regions in the lower parts of the distribution exhibit hysteresis, consisting of long-lived adjustments in employment and labor productivity in response to the shocks. As a consequence, policy tightening aggravates regional inequality and policy easing mitigates it. Finally we provide a structural interpretation of our results using a New Keynesian Currency Union Model with hysteresis effects.

Keywords: Monetary Policy; Regional Heterogeneity; Local Projections; Quantile Regressions; Endogenous Technological Change; New Keynesian

JEL Classification: C32, E32, E52

Non-technical summary

This paper provides a novel perspective on the regional patterns of monetary policy transmission. Using geographically disaggregated data on economic activity in Europe, we show that the output response to short-term interest rate shocks is significantly more pronounced and persistent in poorer than in richer cities and counties. Moreover, while GDP in the upper part of the distribution returns to its pre-shock level after four to five years, the response in the lower parts does not reverse over this period, thus pointing to pronounced hysteresis in output. The heterogeneous incidence of hysteresis in turn implies that monetary policy has a long-lasting impact on regional inequality, with tightening shocks aggravating and easing shocks mitigating it.

In terms of anatomy, we find hysteresis to originate from long-lived adjustments in both employment and labor productivity. At the same time, employment hysteresis is more pronounced and more broad-based across the distribution. In fact, it even extends to the sample mean, as opposed to productivity hysteresis which concentrates in the lower tails of the distribution. As such, our findings confirm labor markets as an important source of hysteresis in the European context. From a theoretical perspective, we show that our main findings are consistent with a New Keynesian Currency Union model in which technological progress is endogenous to the regional unemployment rate.

Our paper points to the use of geographically disaggregated data as a promising avenue for further insights into how exactly monetary policy shocks propagate to the economy. First, by resorting to information on economic activity at a more granular level than that entering the central bank reaction function, it offers a novel strategy to identify exogenous changes in monetary policy. Second, by providing empirical estimates for the monetary policy impact on regional inequality, it closes an important gap in the large and growing literature on heterogeneity in monetary policy transmission.

From a policy perspective, our findings underscore the challenges of calibrating monetary policy in heterogeneous economies. In the euro area context, the debate has typically interpreted these challenges as a cross-country phenomenon. As such, they attracted particular attention during the euro area sovereign debt crisis from 2010-12, which was marked by strong

cross-country divergence in economic performance and raised concerns as to the suitability of a given aggregate monetary policy stance for individual countries. However, our analysis demonstrates that the issue runs deeper: interregional heterogeneity becomes more accentuated at more granular geographical levels and this heterogeneity in turn profoundly alters the implications of a given monetary policy stance in different parts of the economy. These implications emerge as particularly relevant in view of our finding that monetary policy exerts durable impacts on output and employment – a finding that contrasts with the common notion of stabilization policies merely smoothing out fluctuations in these variables around some natural levels.

1. Introduction

There is an increasingly active debate on whether monetary policy tends to dampen or reinforce economic inequality. Much of this debate has centred on the impact of monetary policy on income, consumption and wealth inequality at the household level (Coibion et al., 2017; Kaplan et al., 2018; Auclert, 2019; Bielecki et al., 2021). An aspect that instead has received little attention so far is that economic inequality typically has a pronounced geographical dimension. In particular, some subnational regions generate significantly and persistently higher per capita incomes than others (Jacobs, 1969; Krugman, 1998); and, in many countries, this aspect of inequality has either intensified over recent decades or halted its previous declining trend.¹

A natural question is whether these trends have been influenced by the macroeconomic policy mix, within which monetary policy has played a particularly dominant role since the Global Financial Crisis in most economies. However, empirical evidence on the link between monetary policy and regional inequality is sparse and, *a priori*, it is ambiguous whether monetary policy shocks would reduce or aggravate regional inequality.²

Against this background, the current paper examines this question based on granular data

¹See Ganong and Shoag (2017) and Austin, Glaeser and Summers (2018), Iammarino, Rodriguez-Pose and Storper (2018), and Martin et al. (2018) for recent evidence on the United States, the euro area, and the UK, respectively.

²Some papers find that economic agents with weak balance sheets are more sensitive to monetary policy. This includes evidence on households (see citations above), firms (Kashyap et al., 1994; Gertler and Gilchrist, 1994), and banks (Peek and Rosengren, 1995; Kashyap and Stein, 2000; Altavilla et al., 2020). To the extent that more vulnerable economic agents cluster geographically, these mechanisms may translate into a stronger responses to monetary policy in less prosperous regions. But, other papers highlight channels that induce lower responsiveness in the weaker parts of the respective distributions for firms (Ottonello and Winberry, 2020) and banks (Boeckx et al., 2017). Further, Dolado et al. (2021) find that monetary policy easing raises labour income inequality between high-skilled and low-skilled workers.

on economic activity at the city- and county-level in the euro area and proposes a theoretical mechanism to rationalize our empirical findings.

The use of city- and county-level data, as the lowest aggregation level at which official GDP statistics are available, has two key advantages for this type of analysis. The first pertains to the identification of exogenous variation in monetary policy, which is a key challenge at the heart of empirical monetary economics: as central banks are equipped with explicit macroeconomic stabilization mandates, monetary policy is by construction endogenous to the state and prospects of the economy. As a consequence, it is inherently difficult to disentangle the cause and effect of observed co-movements between policy indicators and macroeconomic aggregates. While the literature has proposed a host of strategies to solve this identification problem, consensus on the actual impact of monetary policy shocks on the economy has remained elusive. For instance, [Ramey \(2016\)](#) documents the wide dispersion in the peak and persistence of the estimated output response to monetary policy shocks in the US.

The regional disaggregation of economic data offers a promising avenue to expand the set of strategies available for solving this identification problem. In particular, our identification strategy allows us to make use of the fact that the mandate of the ECB refers to the euro area aggregate level and the ECB's Governing Council has, since its onset, emphasized that: *“[its] single monetary policy will adopt a euro area-wide perspective; it will not react to specific regional or national developments”* (ECB Governing Council Press Release, 13 October 1998). This euro-area wide focus of monetary policy implies that, controlling for aggregate conditions, variation in short-term interest rates is likely to be exogenous to GDP at the city- and county-level. Hence, by exploiting economic information at a more granular level than that entering central bank reaction functions, we are able to address the risk of simultaneity bias in estimating the real effects of monetary policy; and our main conclusions remain intact when relaxing

the assumption that monetary policy conduct is independent of regional economic disparities, as suggested by [Coibion and Goldstein \(2012\)](#). Similar identification strategies based on sub-national data have been put to productive use in estimating fiscal multipliers (see e.g. [Auerbach and Gorodnichenko \(2012\)](#), [Clemens and Miran \(2012\)](#), [Nakamura and Steinsson \(2014\)](#), [Corbi, Papaioannou and Surico \(2019\)](#), [Guren et al. \(2020\)](#), and [Chodorow-Reich \(2019a\)](#) for a review). But they have received little attention in the monetary economics literature so far.

The second advantage of using these disaggregated data derives from the particularly pronounced disparities arising at the local level. For instance, the variation of per capita GDP at the city- and county-level is around 40% higher than that between countries in the euro area (see [Section 3](#) for details). Accordingly, the disaggregated data provide for a much richer information set to be exploited in analysing the regional transmission of monetary policy. And it allows us to assess whether monetary policy tends to mitigate and accentuate interregional income disparities – an important question in view of the increased scrutiny that the impact of monetary policy on economic inequality has received in public discourse and the increased recognition that rising inequality in many advanced economies exhibits a pronounced interregional dimension ([Coeuré, 2018](#)).

To estimate the dynamic impact of exogenous changes in monetary policy across the entire regional GDP distribution, we combine [Jordà \(2005\)](#)'s local projections method with two complementary approaches for modeling heterogeneity. The first consists of quantile estimation techniques and is based on the novel approach proposed by [Machado and Silva \(2019\)](#) which, building on [Chernozhukov and Hansen \(2008\)](#), provides an avenue to overcome the well-known problems of standard quantile regressions in the presence of fixed effects. The second consists of sub-sample analyses, as suggested e.g. by [Crouzet and Mehrotra \(2020\)](#) in the context of firm size classes, and is based on a break-down of the regional distribution into separate quantile-

ranges. In terms of policy indicators, we focus on standard monetary policy, implemented via policy-controlled short-term interest rates, since this type of policy dominated over our sample period. We do however confirm the robustness of our findings to the use of shadow rates that also capture the impact of non-standard monetary policy on financial conditions. The data are based on Eurostat’s Nomenclature on Territorial Units for Statistics (NUTS) and broken down to the most disaggregated geographical level at which information on economic activity is available (NUTS3).³

Our estimates document a significant and relevant impact of changes in short-term interest rates on regional GDP that strongly intensifies towards the lower end of the distribution. At the sample mean, the response closely resembles the patterns typically found in the literature. After an initial transmission lag, the interest rate coefficient turns significant at a one-year horizon. The impact then further builds up and peaks at the two-year horizon, with the estimates implying a decline in average regional GDP of around 2% in response to a 100 basis point exogenous interest rate increase, before fading out over the remainder of the projection horizon. Moving beyond the sample mean, the quantile estimations point to pronounced heterogeneity in the regional patterns of monetary policy transmission. In particular, output responds more strongly in regions with low versus high per capita GDP and this difference becomes particularly pronounced in the tails of the distribution. Further, the same qualitative patterns emerge when, instead of total GDP, we focus on the capital- and labor-intensive sub-sectors of the regional

³The subnational administrative setup differs substantially across euro area countries and not all countries have a local layer of government equivalent to what is referred to as a “county” for instance in the United States or in Germany. As we discuss in greater detail in Section 3, the NUTS classification tackles this issue by creating a system of economically coherent regions following harmonized standards across countries. For ease of exposition, we refer to NUTS3 regions as the “city- and county-level” throughout this paper but note that this wording is approximate and needs to be interpreted in conjunction with the more precise definitions presented in Section 3.

economies while, among these sub-sectors, capital-intensive production is more responsive, in line with standard notions of the interest rate channel of monetary policy.

The most striking feature of our results is that substantial parts of the regional distribution experience long-lived effects of monetary policy shocks on output. In particular, while GDP in the upper part of the distribution returns to its pre-shock level after three to five years, the response in the lower parts does not reverse over this period, thus pointing to pronounced hysteresis in output. The heterogeneous incidence of hysteresis in turn implies that monetary policy has a long-lasting impact on regional inequality, with tightening shocks aggravating and easing shocks mitigating it.

In terms of anatomy, we find hysteresis to originate from long-lived adjustments in both employment and labor productivity. At the same time, employment hysteresis is more pronounced and more broad-based across the distribution. In fact, it even extends to the sample mean, as opposed to productivity hysteresis which concentrates in the lower tails of the distribution. As such, our findings confirm labor markets as an important source of hysteresis in the European context. And they add to a growing literature arguing that monetary policy may exert durable impacts on output and employment – in contrast to the common notion of stabilization policies merely smoothing out fluctuations in these variables around some natural levels ([Blanchard, 2018](#); [Dupraz, Nakamura and Steinsson, 2019](#); [Jordà, Singh and Taylor, 2020](#); [Ravn and Sterk, 2020](#)). Consistent with recent literature, we find hysteresis-prone regions to be characterized by weaker labour markets, a higher share of employment in small firms, lower business dynamism, and a weaker capacity for innovation.

Our findings add a novel perspective to the literature on the geographical incidence of monetary policy. [Corsetti et al. \(2021\)](#) show that the ECB's monetary policy transmits unevenly across euro area economies, whereas [Carlino and DeFina \(1998\)](#) document heterogeneity in

monetary policy transmission across (clusters of) US states.⁴ At the same time, and notwithstanding the aforementioned benefits, only few papers have studied monetary policy transmission at a richer *regional* disaggregation level. Important exceptions include [Francis, Owyang and Sekhposyan \(2012\)](#) who estimate city-level responses to monetary policy shocks in the US, as well as [Fratantoni and Schuh \(2003\)](#), [Di Maggio et al. \(2017\)](#), [Beraja et al. \(2018\)](#), and [Eichenbaum, Rebelo and Wong \(2018\)](#) who examine the transmission of monetary policy via mortgage markets at similar aggregation levels as the current paper. Our paper is the first to model the impact of monetary policy across the entire regional GDP distribution of a major advanced economy at the city- and county level; to explicitly tackle the question of how the response differs in richer and poorer places; and thus to shed light on how monetary policy affects regional inequality.

Our final contribution is to provide a structural interpretation of our empirical findings. We start from a stylized classical model with a hysteresis channel in which unemployment has endogenous effects on total factor productivity. In this model, we derive analytical conditions under which low-income regions suffer more persistent fluctuations than high-income regions in response to shocks. We then embed this mechanism in a New Keynesian Currency Union model with heterogeneity across regions and analyze numerically the response to monetary policy shocks. Our model is able to qualitatively replicate our empirical findings.

The remainder of the paper proceeds as follows. The next section presents the empirical model and identification strategy. [Section 3](#) describes the data and highlights some salient stylized facts regarding economic activity at the regional level. [Section 4](#) presents our baseline results, complemented by a range of robustness checks in [Appendix B](#). [Section 5](#) provides a

⁴For further literature on the US, see *e.g.* [Di Giacinto \(2003\)](#); [Owyang and Wall \(2009\)](#); [Beckworth \(2010\)](#); [Furceri et al. \(2019\)](#); [Leahy and Thapar \(2019\)](#). For literature studying similar geographical units in Europe, see *e.g.* [Arnold \(2001\)](#); [Arnold and Vrugt \(2002\)](#); [Rodriguez-Fuentes and Dow \(2003\)](#); [Dow and Montagnoli \(2007\)](#).

structural interpretation of our empirical results. Section 6 concludes.

2. Empirical setup

2.1. Baseline model and identification

To study the dynamic effects of monetary policy on regional output, we apply Jordà (2005)'s local projections method consisting in a set of regressions of the form:

$$y_{i,t+h} = \alpha_i + \beta_h i_t + \gamma_h \mathbf{X}_{i,t} + \delta_h \mathbf{X}_{j,t} + \theta_h \mathbf{X}_{k,t} + \varepsilon_{i,t+h} \quad (1)$$

where the dependent variable $y_{i,t+h}$ denotes real GDP in jurisdiction i and year $t+h$; α_i is a set of region-fixed effects; i_t is the monetary policy-controlled short-term interest rate in year t ; $\mathbf{X}_{i,t}$, $\mathbf{X}_{j,t}$, and $\mathbf{X}_{k,t}$ are vectors of time-variant control variables at the local-, country- and euro area-level, respectively; and $\varepsilon_{i,t+h}$ is an error term. All variables enter equation 1 in log-levels, except for the short-term interest rate which enters as a percentage per annum. In the most parsimonious version of the model, the local control variables include the population of each region; the country controls include GDP of the country j in which region i is located; and the euro area controls include GDP and HICP at the euro area level. In the below analysis, we augment this baseline model with a host of further covariates to test the robustness of our findings (see Appendix B). Further, we test whether our main results carry over to the use of alternative monetary policy indicators, including a set of 'shadow interest rates' to capture the impact of non-standard measures affecting the policy stance above and beyond the effect of observed short-term interest rate changes, which may be of particular relevance for the final years of data.

Our main interest is in the impulse response of local output to a change in the short-term interest rate in year t , as captured by the coefficient β_h for each horizon h . The interpretation

of the β_h -coefficients as the causal effect of short-term interest rate changes on local output at horizon h relies on the identifying assumption that, controlling for macroeconomic conditions, monetary policy does not respond to economic activity at the city- and county-level.

This assumption is supported by the following considerations. As highlighted in Section 1, the ECB's monetary policy mandate, with its primary objective to maintain price stability, refers to the euro area as a whole. In keeping with this euro-area wide mandate, the ECB's Governing Council has formally defined its price stability objective in terms of the aggregate euro area inflation rate. Vice versa, it has explicitly ruled out regional developments as a determinant of its monetary policy conduct (ECB, 1998). Moreover, given the dearth and long lags in the availability of regionally disaggregated data it is not in a position to factor in that information at a policy-making frequency.⁵ Accordingly, by controlling for euro area aggregate inflation and activity in the ECB's reaction function, we are able to partial out the variation in short-term rates that does not reflect the systematic central bank response to the economy. Based on this remaining variation, which effectively amounts to using Taylor-rule residuals to identify exogenous changes in policy-controlled short-term interest rates, we can then estimate the impact of monetary policy shocks on regional GDP. As such, the granular data on economic activity helps us overcome the risk of simultaneity bias, which constitutes one of the key challenges in estimating monetary policy effects on the economy.⁶

The remaining explanatory variables serve to further sharpen this identification approach.

⁵In particular, information on economic activity at the NUTS3 level usually becomes available only around two to three years after the period they refer to. This very long lag implies that, at the time policy rates are being set, even the raw data that would allow decision makers to consider regional economic developments are missing.

⁶The assumption that monetary policy responds to aggregate but not to regional shocks also features prominently in the recent model-based analysis of the aggregate implications of regional business cycles by [Beraja et al. \(2019\)](#), as well as in the conceptual discussion of the increasingly widespread use of regional data in macroeconomics by [Chodorow-Reich \(2019b\)](#).

The inclusion of country-level GDP further severs the link between the monetary policy reaction function and regional GDP as the dependent variable. The population variable accounts for the large cross-sectional heterogeneity in the size of the regions.⁷ And the region-fixed effects allow us to control for a host of other, unobserved, factors that may confound the causal interpretation of our estimates. Moreover, we subject our baseline specification to a range of robustness tests ([Appendix B](#)). In the spirit of [Romer and Romer \(2004\)](#), these include the use of real-time ECB forecasts as further macro controls, as well as alternative indicators of the monetary policy stance and a broader set of covariates. Following [Coibion and Goldstein \(2012\)](#), we also present robustness checks in which we relax the identifying assumption that monetary policy conduct is independent of regional economic disparities.

To assess the response of regional economic activity to monetary policy shocks at the sample mean, we estimate equation 1 via ordinary least squares (OLS). Inference is based on [Driscoll and Kraay \(1998\)](#) standard errors that account for cross-sectional and temporal dependencies in the data. In terms of lag-length, we follow the heuristic from the first step of the [Newey and West \(1994\)](#) plug-in procedure, which in our data set implies two lags (see [Hoechle \(2007\)](#) for details). When experimenting with higher lag-lengths, however, we found the estimated standard errors to remain largely unaffected.

2.2. *Modeling heterogeneity*

Besides estimating the mean response of regional GDP to monetary policy shocks, our aim is to also shed light on how this response differs across the (conditional) GDP distribution. To this end, we rely on two complementary approaches, the first consisting of quantile regressions

⁷An alternative would be to directly specify the dependent variables in per-capita terms. Compared to this approach, our choice of specification is somewhat more flexible as it does not restrict the coefficient on the population variable to be equal to one. At the same time, when experimenting with log per-capita GDP as dependent variable, we obtained very similar results.

and the second of sub-sample analysis.

As regards quantile estimation, the seminal approach proposed by [Koenker and Bassett \(1978\)](#) in principle provides a flexible way to model the conditional distribution of the dependent variable. At the same time, it is not ideally suited for panel data models with fixed effects when the cross-section is large relative to the time dimension, in which case the estimates are prone to incidental parameter problems (see, e.g., [Lancaster, 2000](#)).⁸ To address this issue, we therefore employ the quantile regression approach recently proposed by [Machado and Silva \(2019\)](#), which enables us to control for unobserved heterogeneity while estimating quantile-specific coefficients of the covariates in our model via location- and scale-functions:⁹

$$y_{i,t+h} = \alpha_i + \beta_h i_t + \gamma_h X_{i,t} + \delta_h X_{j,t} + \theta_h X_{k,t} + (\lambda_i + X_{i,t}\mu)\varepsilon_{i,t+h} \quad (2)$$

$$Q_y(\tau|X_{i,t}) = (\alpha_i + \lambda_i q(\tau)) + \beta_h i_t + \gamma_h X_{i,t} + \delta_h X_{j,t} + \theta_h X_{k,t} + X_{i,t}\mu q(\tau) \quad (3)$$

⁸A related issue arises with regard to the interpretation of the coefficients. In our application, the regions populating the upper and lower parts of the GDP distributions *conditional on the fixed effects* may be very different than those in the respective unconditional distributions. For instance, it is possible for relatively prosperous regions to temporarily end up in the bottom part of the distribution in years in which their GDP exhibits a transitory drop relative to its sample mean. However, our intention is to sort the cross-sectional units according to the more persistent aspect of regional inequality, which requires an alternative approach.

⁹In the quantile regressions, we again account for potential error correlation across space and time. To this end, we resort to two-way clustering, given the Driscoll-Kraay correction used for the mean regressions is not available for the quantile estimator. As suggested by [Machado and Silva \(2019\)](#), we also tested whether our conclusions may be biased due to the large ratio of cross-sectional units relative to time periods via the split-sample jackknife estimator developed by [Dhaene and Jochmans \(2015\)](#); this also left our results qualitatively unaffected and, if anything, led to somewhat more accentuated heterogeneity in the responses across regions.

Overall, the choice of methodology places our paper in a growing literature using quantile estimation techniques in macroeconomic applications, as recently popularized by [Adrian, Boyarchenko and Giannone \(2019\)](#) and applied to a local projections model estimated on a country panel by [Adrian, Grinberg, Liang and Malik \(2018\)](#).

For the sub-sample analysis, we group the regions according to their position in the per-capita GDP distribution and allow the coefficients in equation 1 to differ across groups. Specifically, we define a dummy variable D_{it}^d for each decile d that is 1 for all regions i whose per-capita GDP falls within this decile in year t and zero otherwise. We then interact this dummy with all explanatory variables in equation 1, including the fixed effects, to obtain decile-specific coefficient estimates (for a similar approach in the context of firm size distributions, see [Crouzet and Mehrotra \(2020\)](#)). Further, we adopt an analogous approach for other quantiles, such as the interquartile range, in the results presented in section 4.

As pointed out by [Koenker and Hallock \(2001\)](#), this type of sub-sample analysis is conceptually distinct from quantile regressions. However, as we show in section 4, the two methods yield mutually consistent and qualitatively similar conclusions in the application considered in the current paper. Further, besides offering an informative cross-check, the sub-sample analysis provides a more flexible framework to characterize differences in the adjustment to policy shocks across regions. In particular, quantile regressions, as an inherent feature, model heterogeneity in relation to the conditional distribution of the dependent variable. But our aim is to not only study heterogeneity in the response of economic activity but also to understand its anatomy, *inter alia* by testing for differential responses in a broader set of dependent variables, such as productivity and employment. The quantile regressions would only allow us to check for differential responses across the distribution for the respective choice of alternative dependent variables. The sub-sample analysis instead is amenable to testing whether the response of

these alternative dependent variables also differs across the per-capita GDP distribution.

3. Data and stylized facts

3.1. Sources and definitions

Our analysis relies on a rich dataset of economic and demographic indicators at the subnational level in Europe. The data are based on Eurostat's Nomenclature of Territorial Units for Statistics (NUTS), which is a hierarchical system for dividing up the economic territory of the EU into four levels. The highest level (NUTS0) corresponds to the nation-state and the lowest (NUTS3), which we use in the ensuing analysis, roughly corresponds to the city- and county-level. In this section, we describe the main features of our data, whereas further information on their construction and sources is available in [Appendix A.1](#).

The source of this data is the Annual Regional Database of the European Commission's Directorate General for Regional and Urban Policy (ARDECO), which is maintained and updated by the Joint Research Centre of the European Commission, and is designed to be the regional counterpart to the AMECO database. The NUTS3 data offer the maximum degree of geographical disaggregation for which information on economic activity is available. At the same time, the coverage in terms of economic variables is relatively limited at the NUTS3 level compared to more aggregated data sets.¹⁰

Our main regional variables are gross domestic product (GDP) and gross value added (GVA), both deflated to 2015 Euros, as well as population and employment. Further, we make use of the breakdown of regional GVA into six sectors of the economy, corresponding to the disaggregation in NACE Rev.2 of: agriculture, forestry and fishing; industry less construc-

¹⁰For instance, data on inflation and unemployment are available only at higher regional aggregation levels; see *e.g.*, [Beck, Hubrich and Marcellino \(2009\)](#), [Özyurt and Dees \(2015\)](#), and [Belke, Haskamp and Setzer \(2016\)](#) for recent studies using these variables.

tion; construction; financial and business services; wholesale, retail, transport, accommodation and food services, information and communication; and non-market services. We merge the sectors of industry and construction to capture the capital-intensive part of regional GVA and the sectors of financial and business services, wholesale, retail, transport, accommodation and food services, information and communication and non-market services to capture the labor-intensive part.

We complement this regionally disaggregated information with euro-area and country-level variables for real GDP and HICP from the AMECO database of the European Commission. Following the bulk of the related literature, the 3-month Euribor serves as our measure of the policy-controlled short-term interest rate (see, *e.g.*, [Coenen et al. \(2018\)](#), [Smets and Wouters \(2003\)](#), [Faust et al. \(2003\)](#)). We source this variable from the Area Wide Model (AWM) database. Further, we also rely on the AWM database for information on oil prices (expressed in US Dollar per barrel) and on country-specific long-term sovereign interest rates. The nominal effective exchange rate, considering 67 trading partners, is taken from Bruegel.

The baseline sample includes all NUTS3 regions from the eleven initial euro area member states, excluding Luxembourg, over the period 1999-2014 and from Greece over the period 2001-2014. The starting point of the sample corresponds to the year the euro currency was introduced and the end-point to the last year in which the ECB steered its monetary policy stance primarily via short-term interest rates (see [Rostagno et al. \(2019\)](#) and [Appendix B.4](#)).¹¹ In some parts of our analysis however, we also draw on extended sample periods that either start already prior to the introduction of the euro currency or go through 2018, the last year for

¹¹Luxembourg is excluded because it consists of just one NUTS3 region. The sample for Greece starts in 2001 because this is when the country introduced the euro. We also follow the literature in excluding the five NUTS3 regions of the French overseas territories of Guadeloupe, Martinique, French Guiana, La Réunion and Mayotte, along with the Portuguese autonomous regions of the Azores and Madeira ([Becker et al., 2010](#)).

which NUTS3-level data are currently available; see [Appendix B.4](#) for further detail. Our final baseline sample consists of a panel of 887 NUTS3 regions over 16 years.

3.2. Heterogeneity across space and time

A comparison of per capita GDP distributions at the regional and national level demonstrates how the degree of heterogeneity intensifies with greater disaggregation (see [Table 1](#)).¹² For instance, in 2014, per capita GDP at the country level ranged from €16,235 in Portugal to €45,083 in Ireland.¹³ While substantial, this difference is dwarfed by the dispersion in regional per capita GDP, which ranged from €9,401 for Serres, a region in Northern Greece, to €137,383 in Wolfsburg, a city in the German state of Lower Saxony.

Table 1: Heterogeneity at different aggregation levels

		Mean	St. Dev.	Min	Max	10 th	Median	90 th
GDP	<i>Country</i>	32017	9888	16235	45083	16974	36432	40127
	<i>Region</i>	29843	13235	9401	137384	16708	27803	43064
CI-GVA	<i>Country</i>	7100	3107	2378	12440	3285	7118	10145
	<i>Region</i>	7925	6424	13	91603	2606	6880	13125
LI-GVA	<i>Country</i>	21115	6240	11198	28787	11479	23070	28105
	<i>Region</i>	18334	8571	5235	79901	10945	16463	27406
Observations	<i>Country</i>	11						
	<i>Region</i>	887						

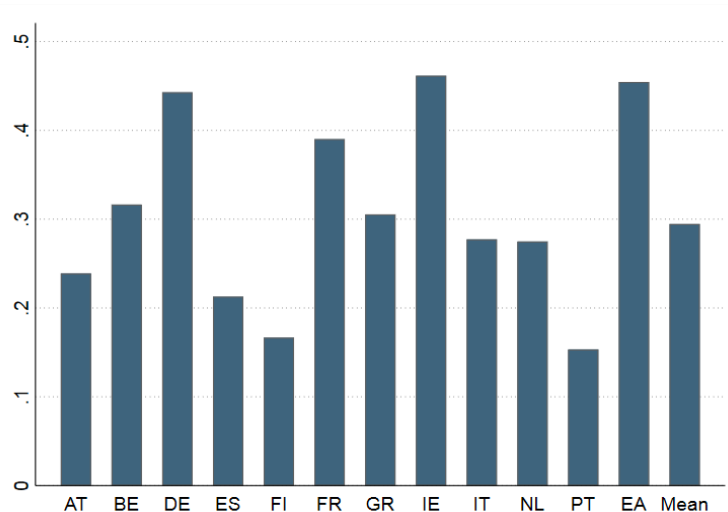
Notes: Figures refer to real GDP and real GVA per capita in 2014 at the NUTS3 (region) or NUTS0 (country) level. Capital-intensive GVA (CI-GVA) is calculated as the ratio of GVA in the sectors of industry and construction over total population. Labor-intensive GVA (LI-GVA) is calculated as the ratio of GVA in the sectors of financial and business services, wholesale, retail, transport, accommodation and food services, information and communication and non-market services over total population.

Moreover, this pattern of greater heterogeneity at more granular aggregation levels is not

¹²Unless otherwise noted, per capita GDP refers to the ratio of real GDP over population throughout the remainder of the paper.

¹³As is well-known, the Irish GDP figures tend to be distorted upwards by the activities of multinational companies. But the difference between the national and regional distribution remains similarly striking when considering the Netherlands, which recorded the second highest per capita GDP in 2014 (of €40,127), as the upper bound for the national distribution.

Figure 1: Coefficient of variation of regional GDP per capita



Notes: The coefficient of variation (CV) is computed as the ratio of the standard deviation to the mean of all NUTS3 regions within each country in 2014, except for: the bar denoted EA, which refers to the CV over all NUTS3 regions in the sample; and the bar denoted Mean, which refers to the unweighted average of the eleven within-country coefficients of variation displayed in the graph.

just confined to a few extreme outliers, but it is a general feature of the respective distributions.

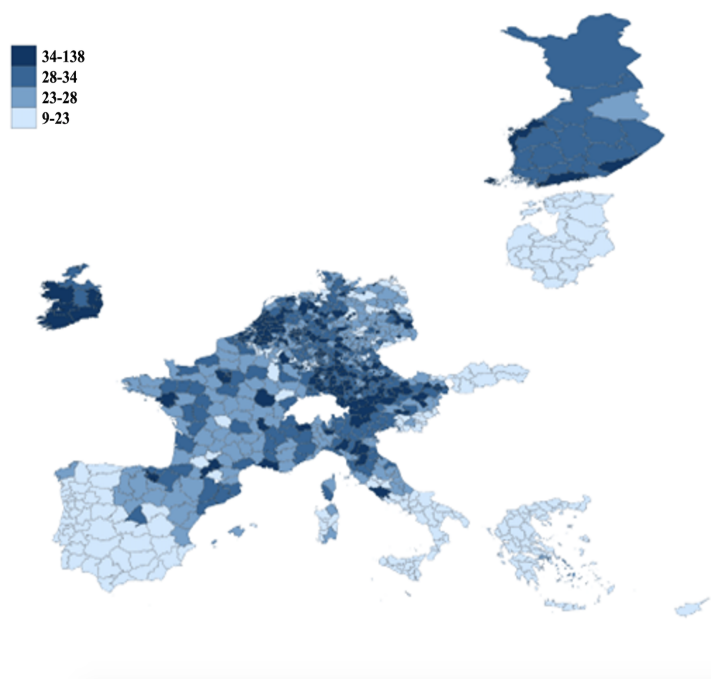
For instance, the coefficient of variation (CV), computed as the ratio of the standard deviation over the mean of per capita GDP, is around 40% higher at the regional than at the national level (see Table 1).

Further, an even more pronounced degree of dispersion emerges for gross value added (GVA), and in particular for the capital-intensive part of GVA, whose CV at the regional level is twice as high as that at the national level. Thus, the dispersion not only in overall activity but also in its composition become more accentuated at the disaggregated level, consistent with the typical patterns of spatial concentration of different types of productive activity (Krugman, 1991; Glaeser et al., 1995).

Importantly, these disparities also arise within the countries. In fact, within-country dispersion in per capita GDP in some cases reaches similar levels as for the euro area as a whole; and the average within-country dispersion in 2014 was almost two thirds of that observed for the euro area (Figure 1; similar patterns emerge for the country-specific time-series averages as

visible from Figure A.18 in Appendix A.1). Accordingly, the analysis of regional disparities in the euro area is not just another way of looking at cross-country differences and putting them under a magnifying glass; instead, it addresses an important aspect of the unequal geography of economic activity that transcends the well-studied cross-country perspective.

Figure 2: Heterogeneity in GDP per capita



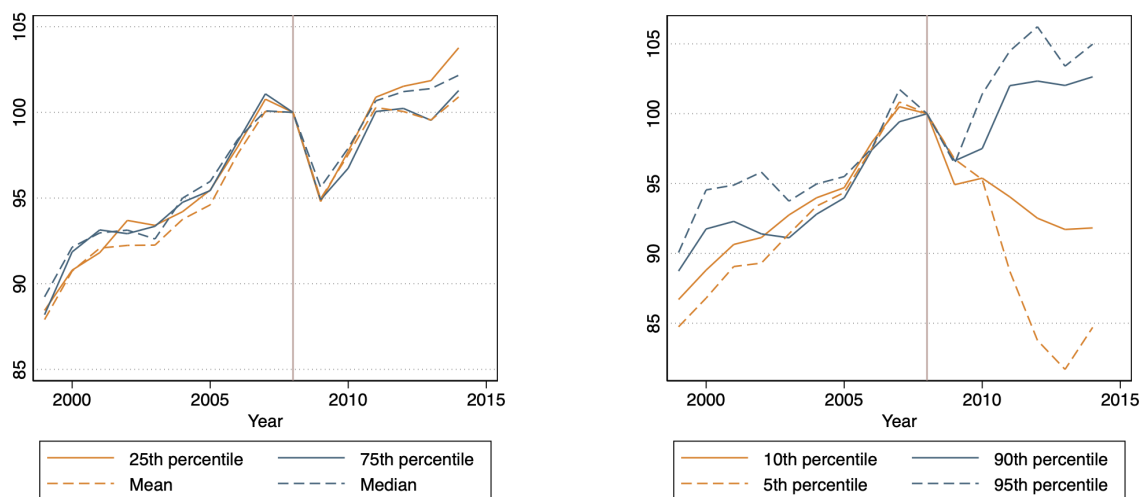
Notes: Figures refer to real GDP per capita in 2014 at the NUTS3 level. Different shadings refer to quartiles.

Moreover, notwithstanding some instances of geographical clustering of more and less prosperous regions, the patterns of within-country inequality are fairly nuanced. For instance, in Italy poorer regions tend to cluster in the South, in Germany in the East, and in Spain in the West (Figure 2). As such, some form of core-periphery divide, which is often emphasized in the context of different countries in the euro area, also exists at the regional level within some of the countries. However, there are also instances of strong disparities arising between regions with a close geographical proximity, as visible from regions in the lower quartiles punctuating clusters of regions in the upper quartiles, and vice versa. Overall, the patterns of heterogeneity within countries are thus fairly diverse, which implies that a less granular subnational disaggregation of the data, such as by provinces or states, would miss important aspects of regional

inequality.

Beyond this static perspective on regional inequality, also its dynamics offer interesting observations, especially in the aftermath of the global financial and economic crisis. This crisis triggered a steep fall in euro area economic activity, followed by a double-dip recession, both of which is clearly visible at the mean of the regional per capita GDP distribution (Figure 3). Similar patterns emerge for the median, whereas the dynamics for other parts of the distribution show a nuanced pattern. While the 25th and 75th percentiles have mostly moved in lockstep, and regions in the 25th percentile even recovered more swiftly than those in the 75th percentile, vastly different trajectories emerge for the outer parts of the distribution. For instance, regions in the 90th percentile on average experienced a solid recovery after the 2009 recession, whereas per capita GDP in the 10th percentile just continued drifting down after the crisis and only showed a mild turn-around in 2014. And this drifting-apart of poorer and richer regions even intensifies when considering the wedge between the 5th and 95th percentile.

Figure 3: Evolution of average per capita GDP in selected percentiles



Notes: The lines show the normalized percentiles of regional GDP per capita in the total sample of EA11. The percentiles have been normalized to 100 in 2008.

Taken together, these stylized facts confirm that the city- and county-level data used in the current paper provide a very interesting setting to study heterogeneity in monetary policy

transmission. Of course, the basic fact that disparities in economic structures and performance become particularly accentuated at this level is well-established in the urban economics literature (Glaeser, Scheinkman and Shleifer, 1995). But it has so far received little attention in the monetary economics literature. The next section closes this gap by presenting evidence on the effects of monetary policy on regional inequality.

4. The heterogeneous impact of monetary policy on regional output

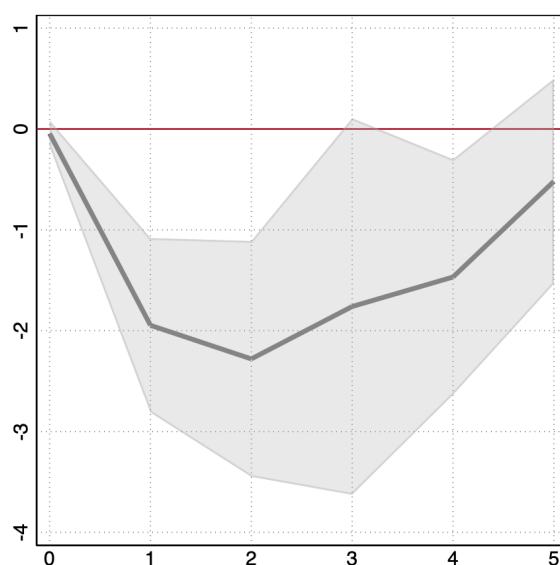
The impulse response functions (IRFs) point to a significant and economically relevant impact of exogenous changes in short-term interest rates on regional GDP; and this impact becomes stronger and more persistent when moving towards the lower parts of the distribution. We next describe these findings in greater detail, while providing a range of robustness checks in [Appendix B](#).

4.1. Monetary policy responses at the mean

As a starting point, and to benchmark our estimates against the related literature, we first focus on the response at the sample mean, estimated via OLS. The following patterns emerge for a 100 basis point hike in short-term interest rates (Figure 4). After an initial transmission lag, the downward impact on economic activity turns significant at a one-year horizon. This impact further builds up and peaks at the two-year horizon, with the impulse response functions implying a contraction of around 2% in regional GDP in year $t + 2$, and then gradually fades out over the remainder of the IRF horizon.

Overall, the estimates closely resemble the familiar patterns for the output response to monetary policy shocks found in the macroeconomic literature for higher levels of aggregation. For instance, the peak effect of a 2% output contraction in response to an exogenous 100 basis point short-term interest rate hike is close to the median of the estimates emerging from prominent

Figure 4: Impact of monetary policy on regional output at the mean



Notes: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands.

contributions to the literature studying the US economy, as reviewed in [Ramey \(2016\)](#) (see Table 1 in that paper). The peak effect is also broadly similar to that deriving from the estimated DSGE model of [Smets and Wouters \(2003\)](#), which often serves as a benchmark for the literature on the euro area economy. Moreover, the two-year transmission lag between the policy rate hike and its peak impact, as well as the gradual, but incomplete, fading out of the output response in subsequent years is comparable to the patterns in [Smets and Wouters \(2003\)](#). This broad match between the estimated responses is reassuring as it allays potential concerns of time aggregation bias arising from the annual frequency of our data ([Marcet, 1991](#); [Hansen and Sargent, 1991](#)).

4.2. *Heterogeneity across quantiles*

Moving beyond the mean, the quantile estimates point to pronounced heterogeneity in the regional patterns of monetary policy transmission. Already for the interquartile range, notable

differences emerge in terms of both, the peak and the persistence of the impacts (Figure 5). Output, in both parts of the distribution, hits its trough in the second year after the shock, but the point estimates point to a somewhat deeper contraction at the lower than at the upper quartile. Moreover, this gap widens over the horizon, as output in the upper quartile rebounds, with the point estimate becoming statistically insignificant in year $t + 5$, whereas output in the lower quartile recovers at a much slower pace and fails to return to its initial level by the end of the horizon.

Moreover, these differences in the output response become markedly more accentuated when considering the outer parts of the distribution. For example, for the 10th percentile the contraction in GDP reaches a trough of -2.5%, compared to -2.0% for the 90th percentile (Figure 6); and the maximum contraction at the 5th percentile, standing at -2.6%, is more than one third deeper than at the 95th percentile (see Figure A.19 in Appendix A.2). Also, the contrast between the quick recovery in the upper part of the distribution and the persistent contractionary effect in the lower part becomes even starker for these percentile pairs: while output recovers to its pre-shock levels by the end of the horizon for the upper percentiles, the contraction does not reverse for the lower ones; for the latter, instead, the point estimates essentially move sideways from horizon $t + 1$ on and remain close to the trough by the end of the horizon. Finally, while the confidence intervals in the interquartile comparison are fairly close together, and over large parts of the horizon overlap, the difference in the outer tails is clearly statistically significant over the later years of the horizon.

Further, these patterns remain intact when relying on sub-sample analysis instead of quantile regressions (Figures 7 and 8, as well as Figure A.20 in Appendix A.2). In particular, both richer and poorer regions experience a contraction in GDP in response to the monetary policy tightening shock, but the maximum impact and its persistence are substantially higher in

Figure 5: Impact of monetary policy on regional output: 25th versus 75th percentile

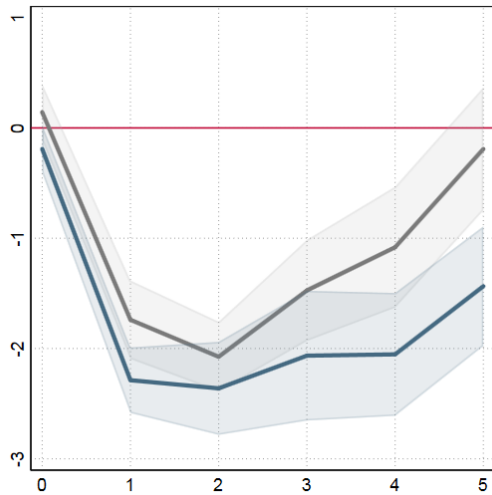


Figure 6: Impact of monetary policy on regional output: 10th versus 90th percentile

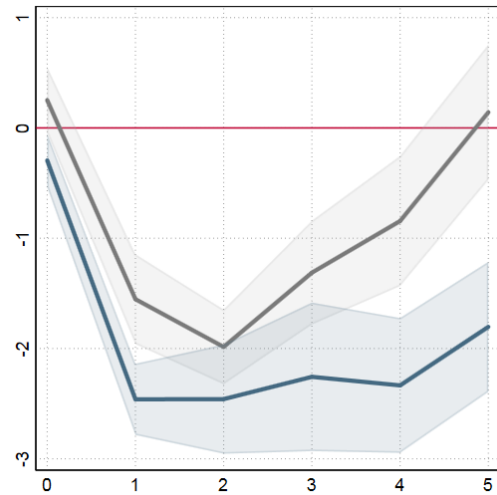


Figure 7: Impact of monetary policy on regional output: upper vs. lower quartile

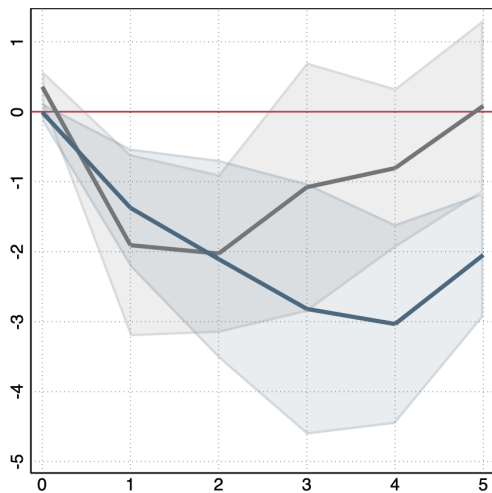
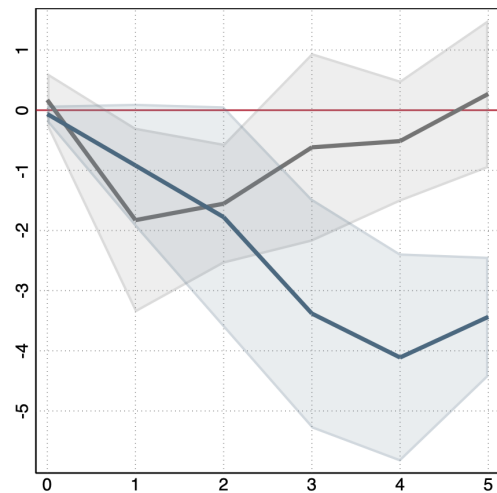


Figure 8: Impact of monetary policy on regional output: bottom versus top decile



Notes: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. Top row refers to estimates from quantile regressions, with grey (blue) lines depicting the estimates for the higher (lower) percentile. Bottom row refers to estimates from the sub-sample analysis, with grey (blue) lines depicting the estimates for the higher (lower) quartile-range.

the latter. Moreover, the estimated extent of contraction in the sub-sample analysis is similar to the quantile regressions, albeit being somewhat more pronounced for poorer regions.¹⁴ As discussed in Section 2.2, this close empirical match between the two approaches is convenient for estimating differential responses across different types of industry (Section 4.3), as well as for exploring the origins of cross-regional differences in the output effects of monetary policy shocks (Section 4.4).

4.3. *The role of industry structures*

Subnational jurisdictions typically differ markedly in their industry mix and a comprehensive literature has documented that the responsiveness of output to monetary policy shocks tends to differ across different types of industry (see, e.g., [Dedola and Lippi \(2005\)](#); [Peersman and Smets \(2005\)](#); and [Carlino and DeFina \(1998\)](#) for representative studies). Our data and research design allow us to explore whether these insights from the related literature are confirmed in our more granular setting and whether regional heterogeneity in industry structures reinforces or mitigates the differential responsiveness to monetary policy shocks.

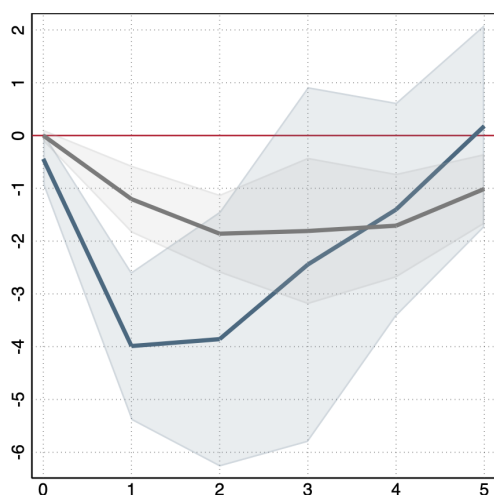
To answer these questions, we again resort to the subdivision of GVA into its capital-intensive and labor-intensive sub-sectors, with the former consisting of construction and industry and the latter of services (see Section 3.2). Based on this breakdown, we conduct two exercises. First, we estimate industry-specific output responses to monetary policy shocks for the sample mean so as to cross-check our priors on the relative interest-rate sensitivity across

¹⁴The resultant higher degree of heterogeneity in impact estimates in the sub-sample analysis relative to the quantile regressions is consistent with the specific features of these different estimation approaches. While the quantile regressions estimate the impact at specific percentiles, the sub-sample analysis estimates average effects within percentile-buckets. Given the impact of policy on regional GDP strengthens in a monotonous fashion, the contraction estimated at the 10th percentile in the quantile regressions tends to be less pronounced than that estimated for the average of the lowest decile as per the sub-sample analysis.

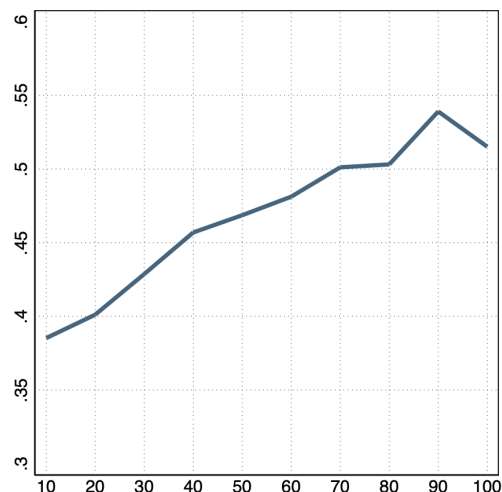
industry-types. Second, we repeat this exercise for the cross-regional GDP distribution via sub-sample analysis so as to test whether differential responses also arise *for a given type of industry*.

Figure 9: Capital-intensive vs. labor-intensive sectors

Impact of monetary policy on regional output across industries



Industry structure across quantiles

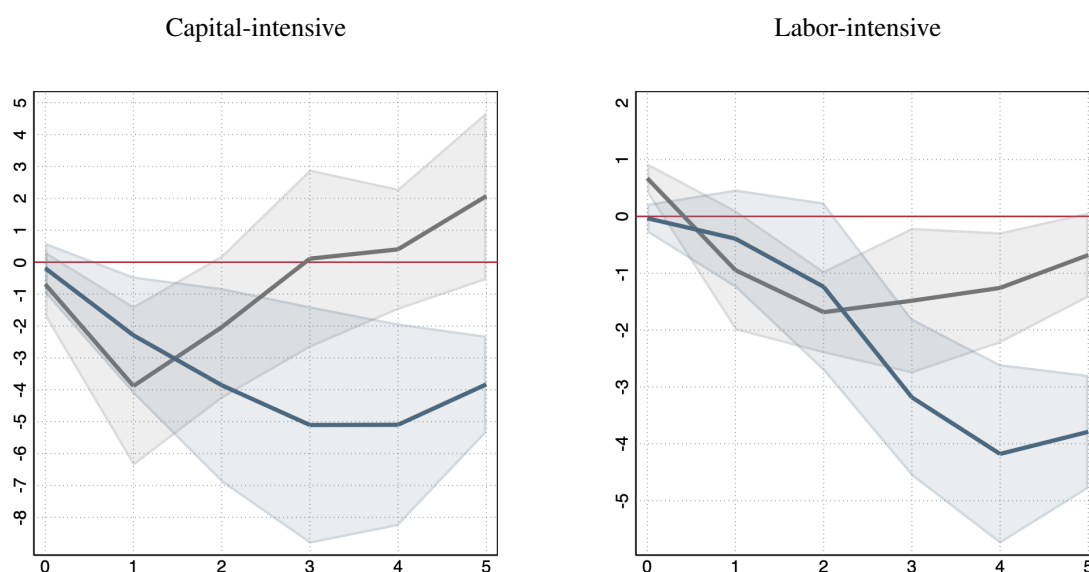


Notes: In the left-hand side (LHS) panel, vertical axis refers to impact of 100 basis point rate hike on regional output in the capital-intensive sector (in blue) and in the labor-intensive output (in grey) at the sample mean (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. In the right-hand side (RHS) panel, vertical axis refers to the ratio of output in the capital-intensive sector over the output in the labor-intensive sector, averaged over the sample period. Horizontal axis refers to quantiles of the regional per-capita GDP distribution.

Consistent with standard notions of the interest rate channel of monetary policy, the impulse responses point to a significantly greater sensitivity of gross value added (GVA) in the capital-intensive sub-sectors (Figure 9, LHS panel). In particular, capital-intensive output contracts more quickly (reaching its trough already in year $t + 1$) and more strongly (with a trough impact roughly double that estimated for labor-intensive output). At the same time, capital-intensive output displays a fairly dynamic V-shaped recovery, whereas labor-intensive output remains near its trough for longer and still remains below its initial level through year $t = 5$, thus pointing to some mild hysteresis, which we study in greater detail below.

A natural question is how the differential responsiveness across industry types interacts with our previous results. For instance, if regions in the lower part of the GDP distribution were to

Figure 10: Impact of monetary policy on regional output: bottom versus top decile



Notes: Vertical axis refers to impact of 100 basis point rate hike (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. The LHS (RHS) chart refers to the GVA of capital-intensive (labor-intensive) sub-sectors. Grey (blue) lines refer to average effects in top (bottom) deciles.

also specialize in capital-intensive production, we could not be sure whether their stronger sensitivity to monetary policy shocks reflects the former or the latter characteristic. However, two pieces of evidence speak against industry structure as an explanation for the greater sensitivity of poorer regions. First, they in fact exhibit a lower share of capital-intensive production than richer regions (Figure 9, RHS panel). Second, even within each type of industry, the differential response across the distribution remains clearly visible: both the capital- and the labor-intensive production contracts more strongly and more persistently in the lower than in the upper part of the distribution (Figure 10); and, within each group of regions, the relative responsiveness across capital- and labor-intensive production is qualitatively similar. Against this background, the inherent differences across more and less prosperous regions also carry over to the industry-breakdown and so does the finding of monetary policy exerting persistent effects on output.

4.4. Sources of hysteresis

The most striking feature of these results is that substantial parts of the regional distribution experience long-lived effects of monetary policy shocks on output. This result clearly contrasts with the common notion of monetary policy as merely causing transitory adjustments in the real economy and, as such, it links to a long-standing debate on potential sources of long-term monetary non-neutrality. This debate has enjoyed a revival in the aftermath of the 2007/2008 financial crisis, which brought renewed urgency to the question whether contractions in aggregate demand may give rise to hysteresis, consisting in lasting declines in the productive capacity of the economy (Yellen, 2016). If so, monetary-policy induced changes in activity may also prove persistent and several recent papers support this conjecture. Most closely related to our paper, Jordà, Singh and Taylor (2020) provide empirical evidence of long-lived effects of monetary policy on output in a panel of advanced economies using a local projections framework similar to that adopted in the current paper.¹⁵

An aspect common to these papers is that they tend to adopt a broad perspective regarding the sources of hysteresis, moving beyond the initial emphasis on the labor market in the seminal contribution by Blanchard and Summers (1986) to also consider productivity and capital accumulation as potential candidates. The debate on which of these potential sources of hysteresis matters (most) is far from settled however. For instance, Jordà, Singh and Taylor (2020) find hysteresis in the impact of monetary policy on the capital stock and on total factor productivity, but not on labor, whereas Blanchard (2018) points to labor markets as an important origin of hysteresis. This question has important implications as to how hysteresis should be modeled and what policy options appear best suited to address it.¹⁶

¹⁵For further empirical evidence see, *e.g.*, Blanchard et al. (2015) and for model-based analysis see, *e.g.*, Reifschneider et al. (2015).

¹⁶For instance, consistent with their emphasis on labor market hysteresis, Blanchard and Summers (1986) focus

Figure 11: Impact of monetary policy on employment and labor productivity in the bottom decile

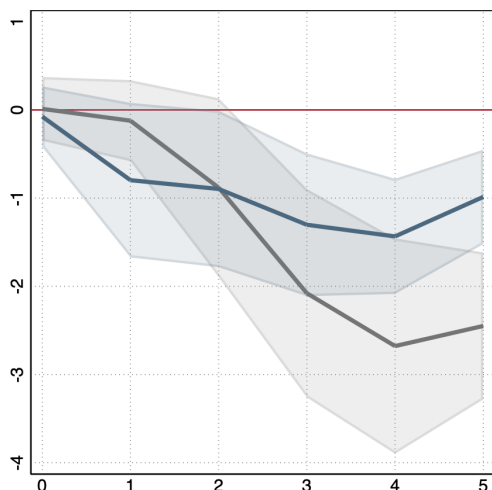


Figure 12: Impact of monetary policy on employment at the mean

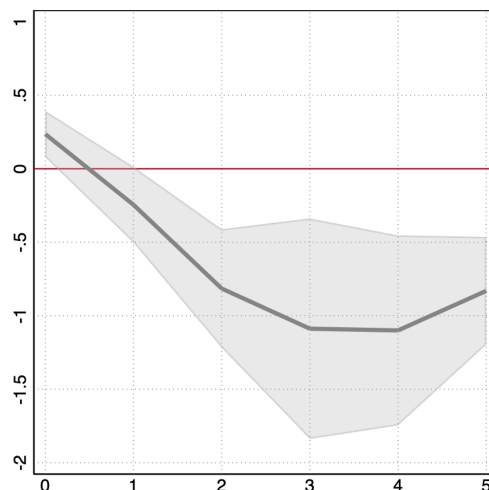


Figure 13: Impact on employment at horizon $h = 5$ across quantile-ranges

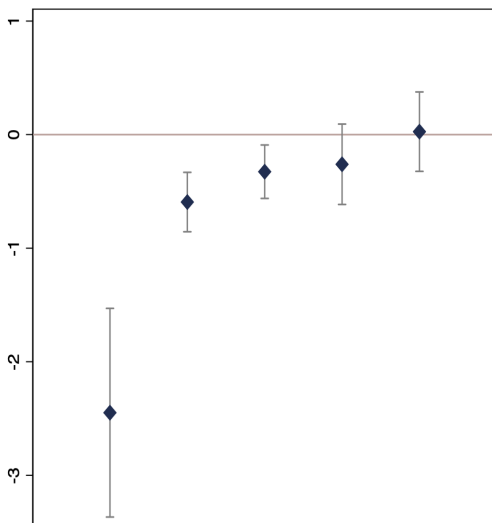
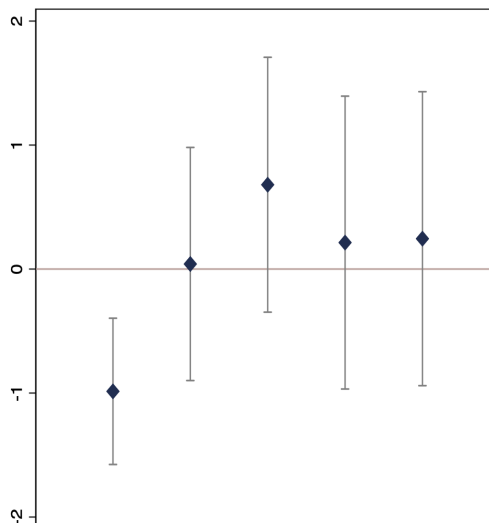


Figure 14: Impact on labor productivity at horizon $h = 5$ across quantile-ranges



Notes: Vertical axis refers to impact of a 100 basis point rate hike on the respective dependent variable (in %), horizontal axis refers to horizon of IRF (in years). In top-LHS panel, the IRFs are for regional employment (in grey) and labor productivity (in blue) in the bottom decile of the per-capita GDP distribution; in top-RHS panel, the IRFs are for regional employment at the sample mean. Solid lines in the top row denote point estimates and shaded areas denote 90% confidence bands. Bottom-LHS panel shows impacts of a 100 basis point rate hike on employment at horizon $h = 5$ (in %) for the decile ranges from: 0-10; 20-30; 50-60; 60-70; and 90-100 (in that order from left to right); bottom-RHS panel shows corresponding impacts on labor productivity. Diamonds indicate point estimates and bars indicate 90% confidence intervals.

Here, we seek to add to this debate by again exploiting the granularity of our data to study the sources of hysteresis. To this end, we re-run our model separately for employment and labor productivity (with the latter being defined as GDP over employment). We first focus on the sub-sample analysis for the bottom decile, which we know from previous exercises to exhibit pronounced hysteresis in the response of output to monetary policy, and then broaden the analysis to further parts of the distribution.

The results point to strong hysteresis in employment, but also labor productivity displays a lasting downward adjustment in response to monetary policy shocks (Figure 11). Employment monotonously falls in response to the shock and, after hitting a trough in year $t + 4$, only marginally recovers to still stand more than 2% below its initial level by the end of the horizon. At the same time, also labor productivity contracts in a persistent manner, albeit less sharply and with a somewhat more pronounced rebound at the end of the horizon. Since our data cover a narrower set of variables than those of [Jordà, Singh and Taylor \(2020\)](#), we cannot disentangle whether the latter finding reflects a lasting erosion of the capital stock or of total factor productivity (TFP). However, to be consistent with the estimated responses of employment and labor productivity, the combined adjustment in the capital stock and TFP has to also consist in a long-lived contraction (conversely, if capital and TFP would remain constant, labor productivity would rise, as the drop in employment would raise the capital intensity).

Further zooming in on employment, the result of strong hysteresis emerges as a fairly broad-based phenomenon, applying to large parts of the distribution. In fact, also at the sample mean employment remains almost 1% below its initial level by the end of the horizon and only shows

on the wage-setting process in modeling hysteresis. [Reifschneider et al. \(2015\)](#) instead also consider hysteresis in capital deepening and multifactor productivity in their extension of the FRB/US model of the Federal Reserve Board staff.

tentative signs of recovery after hitting its trough in the preceding year (Figure 12). Finally, long-lasting employment effects even materialise in the upper half of the distribution, with the coefficient still remaining significant for regions in the sixths decile and only fully reversing in the outer tails (Figure 13). Hysteresis in labor productivity by contrast is less pervasive, as already the third decile displays a full recovery over the horizon (Figure 14).

The uneven incidence of hysteresis, with monetary policy exerting persistent effects in the lower parts and only transitory ones in the upper parts of the distribution, implies that monetary policy may have a long-lasting impact on regional inequality. That hysteresis concentrates in poorer regions appears consistent with their structural features. First, regions in the lower quantiles of the per-capita GDP distribution exhibit a weaker labour market performance, as indicated by their lower employment rates (Figure 15, blue line). Second, poorer regions host a markedly higher share of employees working in small businesses (Figure 15, grey line), and the latter tend to be more exposed to cyclical fluctuations (Crouzet and Mehrotra, 2020). A higher share of fragile firms and low-skilled workers, in turn, implies a stronger vulnerability to policy-induced economic slowdowns (Zens et al., 2020; Duval et al., 2020; Furlanetto et al., 2021). Third, poorer regions are characterised by weaker business dynamism (Figure 16, blue line), which may impede the reallocation of production factors. Policy-induced cutbacks in activity may hence durably weaken the productive capacity of the local economies and thereby translate into a lower GDP trajectory via endogenous growth effects (Moran and Queralto, 2018; Ikeda and Kurozumi, 2019; Anzoategui et al., 2019; Queralto, 2020). Finally, regions in the lower quantiles of the distribution exhibit a weaker innovative capacity, as measured by patent applications (Figure 16, grey line), which may further dampen their resilience to downturns (Martin, 2012).

Figure 15: Employment rates and small business employment across quantiles

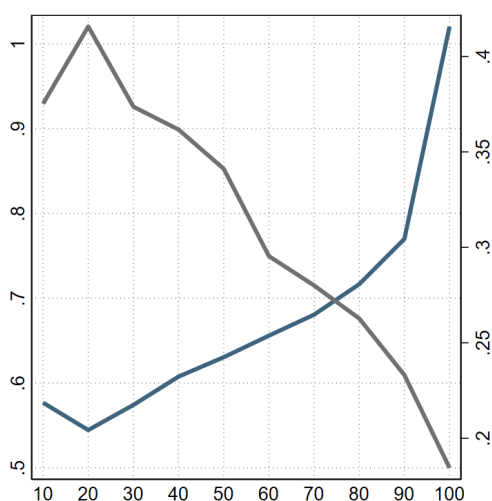
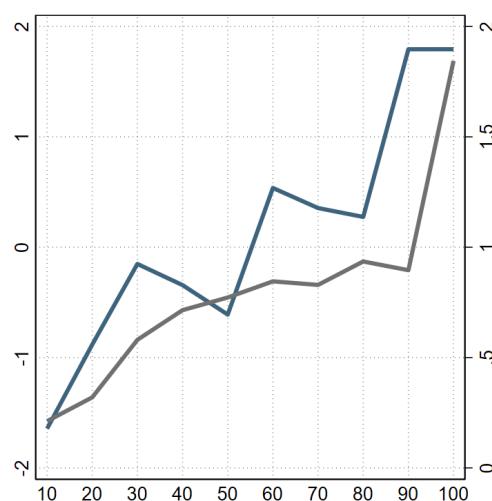


Figure 16: Net business growth and patent applications across quantiles



Source: Eurostat and own computations.

Notes: Horizontal axes refer to quantiles of the regional per-capita GDP distribution. Figure 15 shows average regional employment rates over 2002-2014 defined as the ratio of employed persons to the working age population, i.e. the population aged 15-64 years, (blue line, LHS axis) and the average share of regional employment in small businesses with 1-9 employees for the years 2008-2014 (grey line, RHS axis). Figure 16 shows regional net business growth defined as the birth rate minus the death rate of enterprises on average over 2008-2014 (blue line, LHS axis) and the average number of patent applications per 1000 inhabitants by NUTS-3 region for the years 2001 to 2014 (blue line, RHS axis).

5. Structural interpretation

In this section, we look at our empirical findings through the lens of a stylized model with hysteresis effects. The stylized nature of the model allows us to derive two analytical results that are consistent with the main finding of this paper—namely, that deviations of output from trend are more persistent in Low income regions.

We then embed this stylized model into a New Keynesian model of a Currency Union and show, via numerical simulation, how a common monetary policy shock has differential effects on the persistence of the output, employment, and labour-productivity response across High and Low income regions.

A stylized model Time, t , is discrete and regions are populated by an infinitely lived representative household with preferences over consumption, C_t , and leisure, $1 - N_t$, given by

$$\mathbb{E}_t \sum_{t=0}^{\infty} \beta^t \frac{(C_t)^{1-\sigma}}{1-\sigma} + \xi \frac{(1-N_t)^{1-\varphi}}{1-\varphi}, \text{ where } \beta \in (0, 1) \text{ is the subjective discount factor and } \sigma, \varphi > 0.$$

The budget constraint is given by $P_t C_t + Q_t B_t = B_{t-1} + P_t Y_t$, where P_t is the price of the con-

sumption good and B_t represents the quantity of one-period, nominally riskless discount bonds purchased in period t and maturing in period $t + 1$. Each bond pays one unit of money at maturity and its price is Q_t . Output, Y_t , is produced using the following technology: $Y_t = A_t N_t^{1-\alpha}$, where A_t is Total Factor Productivity (TFP). In equilibrium, $B_t = 0$ and $Y_t = C_t$. In our model, we define unemployment simply as $U_t = 1 - N_t$.

TFP is endogenous and decreasing in the region's unemployment rate, U_t , which individual households take as given. Concretely, we assume the evolution of TFP is given by $A_t = A_{t-1}^\rho (1 - U_t)^\phi$, where $\rho, \phi \in (0, 1)$. Note that ρ captures the strength of the hysteresis effects on TFP. Finally, we impose the parameter restriction $\alpha < \frac{\phi}{1-\rho}$, which ensures that in aggregate, production still features decreasing returns to labour.

The overwhelming majority of hysteresis-type models in the literature establish a link between the endogenous growth engine (for example, capital accumulation, R&D, or learning-by-doing) and the cyclical state of the economy (for example, the level of output, employment, or investment spending). Variations on this theme can be found in [Stadler \(1986\)](#), [King et al. \(1988\)](#), [Engler and Tervala \(2018\)](#), [Bianchi et al. \(2019\)](#), and [Cerra et al. \(2021\)](#). [Chodorow-Reich \(2019a\)](#) and [Cerra et al. \(2023\)](#) provide a comprehensive review of this literature. In our case, we link TFP to the level of unemployment. We make this specific choice for analytical tractability of the model. To demonstrate the generalizability of our results, in [Appendix C](#) we derive very similar analytical results in a Real Business Cycle (RBC) model in which TFP is related the level of region-wide investment spending.

Since we assume prices are fully flexible, the only first-order condition that determines real outcomes is the consumption-leisure trade-off given by $\xi C_t^\sigma (1 - N_t)^{-\varphi} = (1 - \alpha) A_t N_t^{-\alpha}$. Bringing everything together, the equilibrium can be fully described by two endogenous variables: $\{A_t, N_t\}$, and the following two equations: $\xi N_t (1 - N_t)^{-\varphi} = (1 - \alpha) A_t^{1-\sigma} N_t^{(1-\alpha)(1-\sigma)}$

and $A_t = A_{t-1}^\rho N_t^\phi$.

We are now ready to analyze the comparative statics of this economy and present our two key propositions. We begin with our first proposition regarding the effects of hysteresis on the steady state:

Proposition 1 *A sufficient condition for steady state output, Y_t , to be decreasing in the persistence of TFP, ρ , is $\sigma < 1$ (i.e., an intertemporal elasticity of substitution greater than 1).*

See [Appendix C](#) for the proof. The parameter σ measures the strength of the wealth effect of labor supply. When $\sigma < 1$, the substitution effect on labor supply resulting from a lower real wage (e.g., because of a fall in A_t) dominates the positive effect caused by a higher marginal utility of consumption. All else equal, this leads to a decrease in employment. In steady state, the persistence of the TFP process acts as a multiplier to this relationship between labour supply and TFP, resulting in regions with higher TFP persistence experiencing lower incomes.

Our second proposition relates hysteresis to the persistence of output, employment, and labour productivity deviations from trend.

Proposition 2 *The conditional autocorrelation of output, employment, and labour productivity are increasing in ρ (so long as $\sigma < 1$).*

The proof is in [Appendix C](#). However, the key step is to show that the dynamics of these variables can be represented as $ARMA(2, 1)$ processes, from which it is straightforward to derive analytical expressions for their autocorrelation functions. The persistence of output depends on both the persistence of TFP and labour supply. Again, a key requirement is that $\sigma < 1$. This ensures that labour supply is increasing in TFP (and the real wage). If instead $\sigma > 1$, the household compensates for any (endogenous) fall in TFP by raising labour supply, which would cause employment to return back to steady state faster. Thus, even if TFP is more per-

sistent because of a larger ρ value, the labour supply response would cause the effect on the persistence of output to be ambiguous.

The two propositions, taken together, are consistent with the main empirical findings of this paper, which show that deviations of output are more persistent in Low income regions. Moreover, the model and propositions are very general. For example, we do not need to specify the shock that drives the economy away from steady state. Neither do we need to invoke nominal frictions nor any assumptions regarding the “small” open economy nature of a region.

A currency union model Nevertheless, we conclude this section by showing that the above analytical results hold numerically in the richer setting of a New Keynesian Currency Union model. The full model is an extension of [Gali and Monacelli \(2008\)](#), augmented to allow for a non-unitary intertemporal elasticity of substitution and an endogenous TFP process, exactly as described above.¹⁷ The linearized model features region-specific Phillips curves and a union-wide IS curve and monetary policy interest rate rule. We assume all the regions are described by a common set of parameters, except ρ , where half the regions have $\rho = 0$ and the other half have $\rho = 0.925$. An endogenous outcome of the model is that the former (latter) are the High (Low) income regions. Specifically, the calibration implies that steady state income is 25% higher in the High income regions and the unemployment rates are 5% and 7.4%, respectively. The full details of the model derivation and calibration are in [Appendix D](#).

Figure 17 plots the response to an unanticipated union-wide monetary policy tightening. Output and inflation fall in all regions. However, for the High income regions (red-circle line), the fall in output is temporary whereas, for the Low income regions (solid-blue line), output remains persistently below steady state. The differences arise from the endogenous TFP chan-

¹⁷The model is also similar to [Hazell et al. \(2022\)](#) and [Bellifemine et al. \(2022\)](#), which both model regional heterogeneity, although they abstract from endogenous TFP.

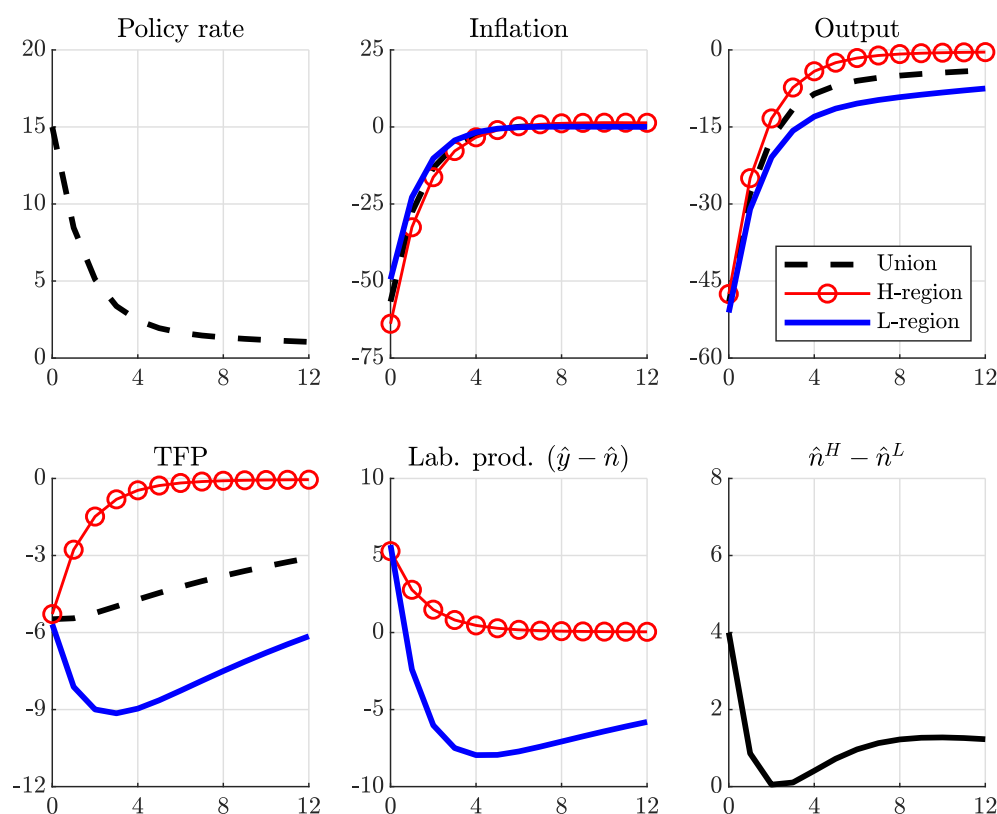
nel. As a result of the fall in employment (rise in unemployment), TFP declines. However, the Low income regions are characterized by hysteresis in that the impact of those declines in TFP become imbedded. In fact, TFP continues to fall for several quarters after impact in the Low income region as the effect of the high unemployment rate on TFP dominates its natural recovery. For the High income regions, TFP instead begins to recover immediately. Like for the output response, employment is further from steady state in the Low income region both on impact and thereafter.¹⁸

While our empirical results do not consider the regional inflation effects of a monetary policy shock, the model predicts that inflation falls less in Low income regions. The intuition for this is to think of the scenario in Figure 17 as the combination of a correlated monetary policy and negative technology shock, with the latter being larger in the Low income region. A negative technology shock lowers output and raises inflation, all else equal, by disrupting the supply side of the economy. This, in turn, pushes up relatively more on inflation in the Low income region.¹⁹

¹⁸The differences in employment outcomes are small across the regions. Hence, the bottom right panel plots the cross-region differences rather than the levels of employment.

¹⁹[Appendix D](#) also contains an analysis of how increasing heterogeneity across regions effects the transmission of monetary policy and the impact on union-wide aggregates.

Figure 17: Response to a monetary policy shock



Note: 1 percentage point monetary policy tightening. Variables are basis point deviations from steady state. The policy rate and inflation are annualized.

6. Conclusion

This paper provides a novel perspective on the regional patterns of monetary policy transmission. Using geographically disaggregated data on economic activity in Europe, we show that the output response to short-term interest rate shocks is significantly more pronounced and persistent in poorer than in richer cities and counties. Moreover, while GDP in the upper part of the distribution returns to its pre-shock level after three to five years, the response in the lower parts does not reverse over this period, thus pointing to pronounced hysteresis in output. The heterogeneous incidence of hysteresis implies that monetary policy has a long-lasting impact on regional inequality, with tightening shocks aggravating and easing shocks mitigating it. In terms of anatomy, we find hysteresis to originate from long-lived adjustments in both employ-

ment and labor productivity. At the same time, employment hysteresis is more pronounced and more broad-based across the distribution. As such, our findings confirm labor markets as an important source of hysteresis in the European context. From a theoretical perspective, we show that our main findings are consistent with a New Keynesian Currency Union model in which technological progress is endogenous to the regional unemployment rate.

Our paper points to the use of geographically disaggregated data as a promising avenue for further insights into how monetary policy shocks propagate to the economy. First, by resorting to information on economic activity at a more granular level than that entering the central bank reaction function, it offers a novel strategy to identify exogenous changes in monetary policy. Second, by providing empirical estimates for the monetary policy impact on regional inequality, it closes an important gap in the large and growing literature on heterogeneity in monetary policy transmission.

From a policy perspective, our findings underscore the challenges of calibrating monetary policy in heterogeneous economies. In the euro area context, the debate has typically interpreted these challenges as a cross-country phenomenon. As such, they attracted particular attention during the euro area sovereign debt crisis from 2010-12, which was marked by strong cross-country divergence in economic performance and raised concerns as to the suitability of a given aggregate monetary policy stance for individual countries. However, our analysis demonstrates that the issue runs deeper: interregional heterogeneity becomes more accentuated at more granular geographical levels and this heterogeneity in turn profoundly alters the implications of a given monetary policy stance in different parts of the economy. These implications emerge as particularly relevant in view of our finding that monetary policy exerts durable impacts on output and employment – a finding that contrasts with the common notion of stabilization policies merely smoothing out fluctuations in these variables around their natural levels.

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Appendices For Online Publication

Appendix A. Data construction and further results

Appendix A.1. Additional detail on the regional data

The regionally disaggregated data are based on a harmonised and integrated breakdown of territorial units established by a European Parliament and Council Regulation (*Regulation (EU) No 549/2013, 2013*). This regulation ensures comparability of classifications across countries and follows two principles. First, it sets a range for the number of inhabitants that each NUTS region should comprise; NUTS3 regions, for instance, should be drawn up in a way that their population falls between 150,000 and 800,000 inhabitants. Second, it relies on existing administrative units, where available; for instance, if the administrative structure of a country includes counties, then these are used to construct NUTS3 regions for municipalities whose population falls below the minimum number of inhabitants; in countries without counties, the NUTS3 regions are created by aggregating smaller administrative units such that they meet the population requirement. The regulation is legally binding, so member states have to comply as a matter of Union law. Moreover, regular revisions to the NUTS classification ensure that any changes in national administrative regions are reflected in the disaggregation without compromising the comparability of subnational indicators.

The source of this data is the Annual Regional Database of the European Commission's Directorate General for Regional and Urban Policy (ARDECO), which is maintained and updated by the Joint Research Centre of the European Commission, and is designed to be the regional counterpart to the AMECO database. The NUTS3 data offer the maximum degree of geographical disaggregation for which information on economic activity is available. The database uses the NUTS 2016 classification, and covers the statistical territory of EU27 and Norway for the period between between 1980, or 1990 for the new member states, and 2015. Nominal mea-

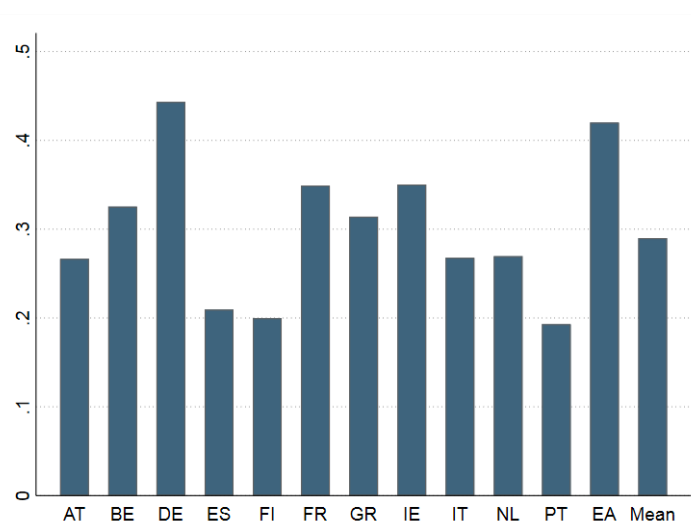
Table A.2: Summary statistics across full sample period (1999 – 2014)

	Mean	Standard deviation	Minimum	Maximum
GDP	28461	11827	9044	137384
Population	352551	469651	8456	6419324
HICP rate	1.87	0.99	-1.69	5.28
Short-term rate	2.36	1.49	0.2	4.64
NEER	107	9.26	88.1	120
Non-oil commodity prices	128	40.0	75.9	201
Oil prices	64.4	33.2	17.7	112
Observations	14088	14088	14088	14088

Notes: All GDP and GVA figures are in real per capita terms. Population is in thousands of people. Non-oil commodity prices are in US dollars and oil prices are UK Brent in US dollars per barrel.

asures of gross value added (GVA), gross domestic product (GDP), compensation of employees, and gross fixed capital formation (GFCF) are deflated to 2015 constant price euros. To construct real GDP at the NUTS3 level, real GVA growth rates at the NUTS2 level are downscaled to the NUTS3 level and then used to deflate GDP at that level. Employment and population are also available at the highest level of regional disaggregation. Sectoral disaggregation refers to the sectors of agriculture, forestry and fishing; industry (excluding construction); wholesale, retail, transport, accommodation and food services; information and communication; financial and business services; and finally, non-market services.

Figure A.18: Coefficient of variation of regional GDP per capita for the full sample period



Notes: The coefficient of variation (CV) is computed as the ratio of the standard deviation to the mean of all NUTS3 regions within each country for the sample period 1999-2014, except for: the bar denoted EA, which refers to the CV over all NUTS3 regions in the sample; and the bar denoted Mean, which refers to the unweighted average of the eleven within-country coefficients of variation displayed in the graph.

The ARDECO Database is based on data from the regional database of Eurostat. The primary advantages of using the ARDECO database over downloading the data directly from Eurostat are two. First, the sample period for the ARDECO database is 1980-2018 for most countries and 1990-2018 for the rest, while for Eurostat regional economic accounts start only in 2000, and a consistent back series of data is not available for the full series. The ARDECO database splices the earlier series to match more recent series. Splicing the data involves using an overlapping period between the European System of Regional and National Accounts (ESA) 79, ESA 95, or ESA 2010 series to extend the ESA 2010 series backwards using ESA 95 and ESA 79 growth rates. The second advantage of the ARDECO database is that they deal with missing data. Even for the period after 2000 that is covered by Eurostat, there is a considerable number of missing values. ARDECO fills in those gaps by scaling up data from sub-regions, extrapolation and interpolation. They additionally implement manual fixes, and scale up the data to AMECO totals. As a result, their final product is a consistent and complete dataset that makes use of and extends upon information provided by Eurostat or AMECO.

Appendix A.2. Additional charts for Section 4

Impact of monetary policy on regional output in the tails of the distribution

Figure A.19: Quantile regressions

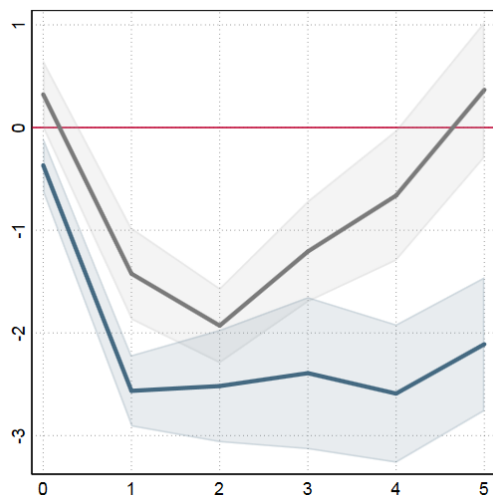
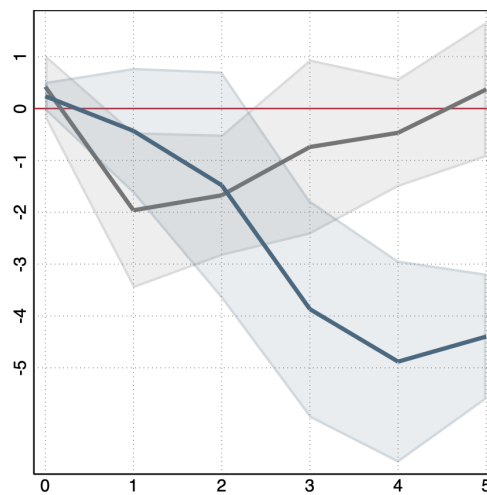


Figure A.20: Sub-sample analysis



Notes: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. LHS-panel shows the 5th percentile (in blue) and the 95th percentile (in grey) of the baseline model estimated via quantile regressions; RHS-panel shows the corresponding average estimates from the sub-sample analysis for the bottom five percent (in blue) and top five percent (in grey) of the per-capita GDP distribution.

Impact of monetary policy during the pre-crisis period (1990-2007)

Figure A.21: Employment

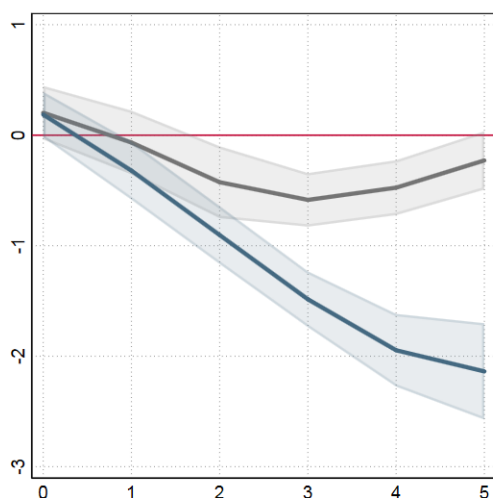
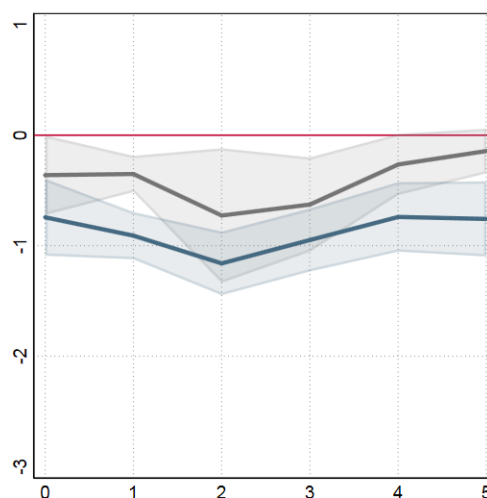


Figure A.22: Productivity



Notes: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. LHS-panel shows the IRFs for employment at the bottom ten percent (in blue) and top ten percent (in grey) of the per-capita GDP distribution during the period 1990-2007. RHS-panel shows the corresponding IRFs for productivity.

Appendix A.3. Construction of forward-looking variables

In one of our robustness checks, we augment the baseline specification with forward-looking variables for GDP growth and HICP inflation (see [Appendix B](#)). These variables are constructed from the quarterly ECB staff macroeconomic projections. These projections are the

natural choice among a variety of alternative sources that provide publicly available macroeconomic projections for the euro area (including for instance other policy institutions or private forecasters). While these projections are a staff-level exercise that does not necessarily have to fully match the Governing Council’s assessment of the economic outlook, they are an integral element of the Governing Council’s information set. This prominent role becomes visible for instance from the fact that the staff projections are reported in the Introductory Statements following the Governing Council’s monetary policy meetings, where they provide the motivation and context for the decisions taken at these meetings.

We choose the last projection year as the reference point because we consider it as the most suitable approximation of the medium-term orientation: while the ECB has clarified that the definition of the “medium-term” may vary depending on the state of the economy and has refrained from defining it in terms of a specific time horizon, existing communication indicates that the outlook for the final projection year carries particular weight in its policy deliberations.

For most of the sample period, the horizon of the ECB projection vintages in year t stretches until year $t + 2$. The ECB changed this convention in December 2016: since then, the forecast horizon for the fourth quarterly vintage of each year has extended to year $t + 3$. For the sake of consistency, and given this change took place only after the end of our baseline sample, we abstract from it in constructing the forward-looking control variables. In summary, we hence calculate the forward-looking inflation variable as:

$$\pi_{t+2|t}^e = \frac{1}{4}(\pi_{t+2|MAR,t}^e + \pi_{t+2|JUN,t}^e + \pi_{t+2|SEP,t}^e + \pi_{t+2|DEC,t}^e)$$

where $\pi_{t+2|t}^e$ is the expected inflation rate for year $t + 2$ entering the central bank information set in year t , $\pi_{t+2|MAR,t}^e$ is the expected annual inflation rate in year $t + 2$ according to the March ECB staff macroeconomic projections of year t , and $\pi_{t+2|JUN,t}^e$, $\pi_{t+2|SEP,t}^e$, and $\pi_{t+2|DEC,t}^e$ are

the corresponding values according to the June, September, and December projections of year t . The forward-looking variable for the rate of real economic growth is calculated in analogous fashion.

Appendix B. Robustness

Appendix B.1. Forecast-based macroeconomic controls

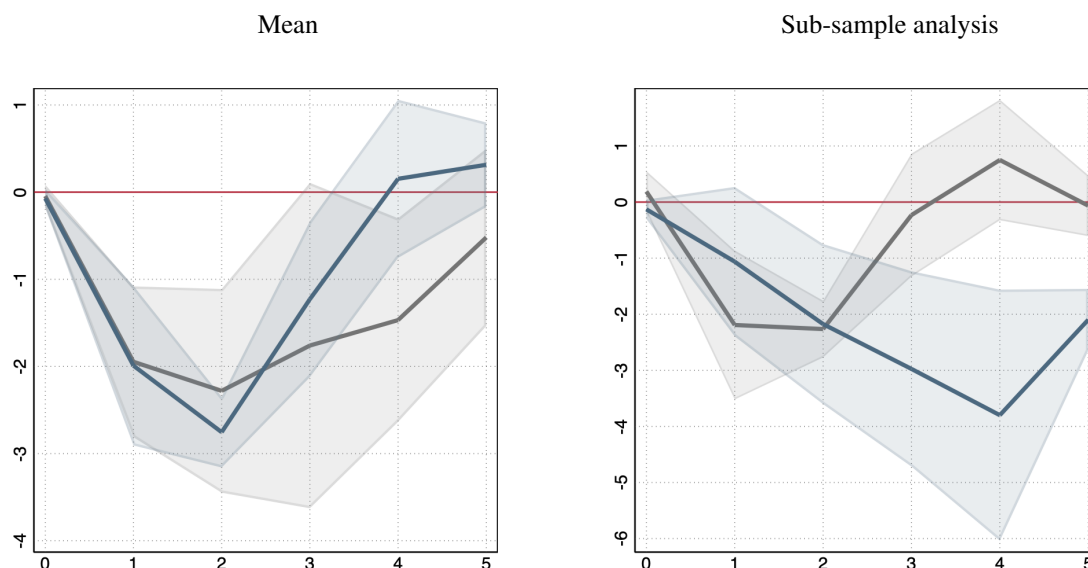
To check the robustness of our main findings, we re-estimate the model with several modifications to our baseline specification. First, we extend the list of controls in equation 1 with forward-looking variables that may shape the ECB’s monetary policy deliberations in real-time. By including contemporaneous measures of aggregate economic activity and prices, our baseline identification strategy essentially amounts to using Taylor-rule residuals as a measure of exogenous changes in policy-controlled interest rates. A potential concern regarding this choice of specification is that the central bank reaction function may depend not (only) on current economic conditions but (also) on their expected evolution over a policy-relevant horizon (Sims, 1992; Romer and Romer, 2004). In fact, the ECB has emphasized that its monetary policy has a “medium-term orientation”, which means that “monetary policy needs to act in a forward-looking manner” (Issing, 1998; ECB, 1998, 2019).

To implement this robustness check, we construct the forward-looking indicators of real GDP growth and HICP inflation from the quarterly ECB staff macroeconomic projections. With HICP inflation, we capture the ECB’s main ‘target variable’ and, by adding GDP growth, we account for the broader economic context of the expected inflation path. As a rule, we use the final year of the respective projection horizon as the relevant reference point and compute the average across the four projection vintages of each year (see Appendix A.3 for additional detail).

The resultant estimates are very similar to our baseline (Figure B.23). For the sample mean, the confidence intervals of the two specifications overlap throughout the IRF horizon; and, for the sub-sample analysis, the alternative specification confirms the weaker and less persistent response of economic activity in the upper than in the lower parts of the distribution. Our

baseline findings thus prove robust to the inclusion of forward-looking control variables for prices and activity.

Figure B.23: Inclusion of expected HICP inflation and GDP growth



Notes: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. LHS-panel shows mean estimates for the baseline (in grey) and the modified model (in blue). RHS-panel shows the corresponding estimates from the sub-sample analysis for the bottom ten percent (in blue) and top ninety percent (in grey) of the per-capita GDP distribution.

Appendix B.2. Controlling for regional dispersion

Our identification strategy rests on the assumption that, after controlling for macro-economic conditions, policy rates are independent of regional developments. As a second robustness check, we test this assumption and assess its relevance for our main conclusions. The analysis is motivated by [Coibion and Goldstein \(2012\)](#) who find that measures of economic dispersion across US states do emerge as relevant determinants of the Federal Reserve's interest rate decisions in estimated Taylor rules.

In applying this approach to our empirical setting, we first regress the policy rate on its own lag, the euro area growth and inflation rates, and the degree of economic dispersion at the regional level (also lagged). Following [Coibion and Goldstein \(2012\)](#), we alternate between three dispersion measures: (i) the population-weighted standard deviation in per-capita GDP;

(ii) the difference between per-capita GDP in the 80th and the 20th percentile of the regional distribution; and (iii) the equivalent measure for the 90th and 10th percentile.²⁰

The regressions provide evidence in favour of [Coibion and Goldstein \(2012\)](#)'s finding that monetary policy rates may respond to subnational economic dispersion. Despite the low number of observations, the weighted standard deviation and 90th to 10th percentile difference carry a significant coefficient (Table B.3).²¹ The negative sign suggests that the policy reaction function indeed embeds a concern for interregional disparities and internalises that these disparities may be mitigated by a more accommodative stance, as documented in Section 4. From an econometric perspective, the results therefore point to the risk of omitted variable bias in our baseline estimations.

We hence extend the baseline equation by including the regional dispersion measures as additional controls and compare the resultant IRFs with our previous estimates. For the sample mean, all three specifications yield a similar timing and trough for the GDP contraction, while exhibiting a somewhat higher persistence and, when including the 90th to 10th percentile difference as a control, some whipsaw pattern in the IRF (Figure B.24). Likewise, the sub-sample analysis again confirms the uneven transmission of monetary policy to poorer and richer regions, albeit again with a somewhat jagged IRF for the specification including the 90th to 10th percentile difference.

²⁰While [Coibion and Goldstein \(2012\)](#) base their dispersion measures on unemployment rates, the use of GDP-based measures appear preferable in our setting because GDP is the main dependent variable of interest also in the other parts of our analysis. In constructing the weighted standard deviation, we also experimented with GDP instead of population-weights, but found the results to be unaffected by this choice. To lower the hurdle for rejecting our H_0 that regional dispersion does not affect policy rate setting, the regressions make use of the full time series dimension of the data, ranging from 1990 to 2018. Since euro area inflation and growth forecasts are unavailable for the pre-euro period, we estimate the policy rule based on realised values for these variables.

²¹The coefficients on inflation and economic activity are also highly significant, have the expected sign, and for inflation are quantitatively similar to the standard Taylor rule prescription of 0.5.

Table B.3: Estimated policy reaction functions including regional economic dispersion measures

	None	Population-weighted S.D.	80 th -20 th Pctl.	90 th -10 th Pctl.
Policy Rate (1 st lag)	0.709*** (0.093)	0.487*** (0.086)	0.593*** (0.083)	0.508*** (0.100)
Inflation	0.651** (0.272)	0.716*** (0.195)	0.696*** (0.213)	0.667*** (0.216)
Growth	0.040** (0.015)	0.029** (0.013)	0.032** (0.014)	0.033** (0.013)
Dispersion (1 st lag)		-0.047*** (0.016)	-0.035 (0.023)	-0.025** (0.009)

Notes: Each column refers to estimates from a separate time series regression over the period from 1990 to 2018, with each including a different regional economic dispersion measure. Newey-West standard errors are reported in parentheses. Statistical significance at the 10%, 5%, and 1% levels is marked by *, **, and ***, respectively.

Appendix B.3. Additional covariates

To guard against other sources of omitted variable bias, another set of robustness checks extends the fairly parsimonious baseline model with a set of factors that may correlate with both, regional activity and policy rates. Prominent examples include global oil prices and exchange rates against major trading partners. These variables routinely feature in ECB policy communication;²² and they appear as potentially important controls at a disaggregated level where the typical spatial concentration of different types of economic activity may render certain regions more and others less responsive to these types of shocks (see, *e.g.* [House, Proebsting and](#)

²²In particular oil price developments are a regular element in the economic assessment communicated via the Introductory Statements to the Governing Council’s monetary policy press conferences. Exchange rates feature less regularly, but do occasionally enter the Introductory Statements, such as on 25 January 2018, when the Governing Council communicated that: “*recent volatility in the exchange rate represents a source of uncertainty which requires monitoring with regard to its possible implications for the medium-term outlook for price stability*”. Moreover, the relevance of exchange rates has, at various occasions, been emphasized in the question-and-answer sessions during these press conferences by ECB Presidents. For instance, President Trichet, on 7 October 2010, pointed out that “*excess volatility and disorderly movements in exchange rates have adverse implications for economic and financial stability*”; and President Draghi, on 25 January 2018, stated that “*exchange rates are important for growth and for price stability*”.

Controlling for regional dispersion in the local projections

Figure B.24: Mean

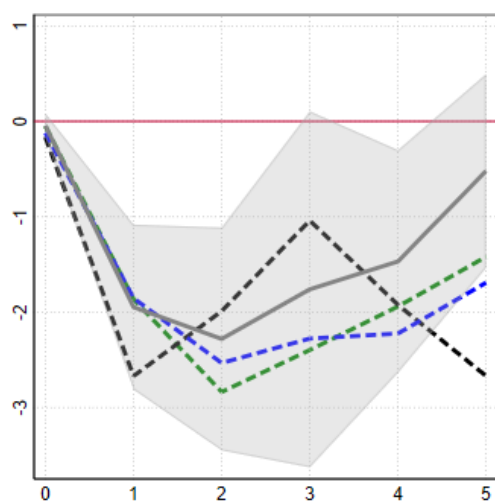


Figure B.25: Sub-sample analysis with population-weighted standard deviation

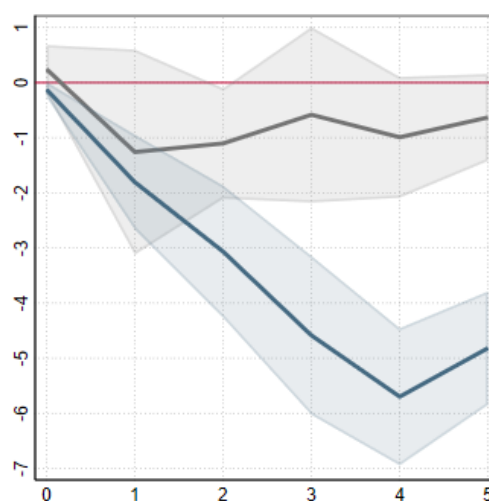


Figure B.26: Sub-sample analysis with 80^{th} - 20^{th} percentile difference

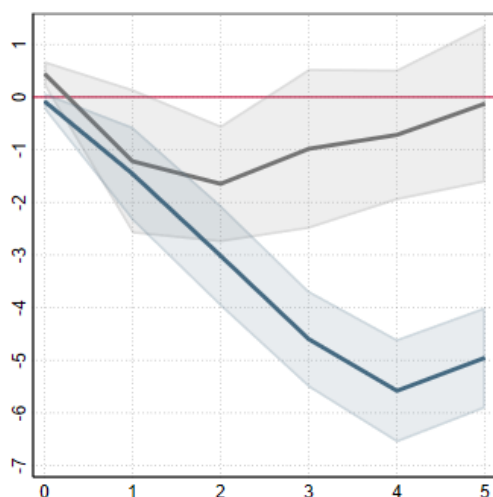
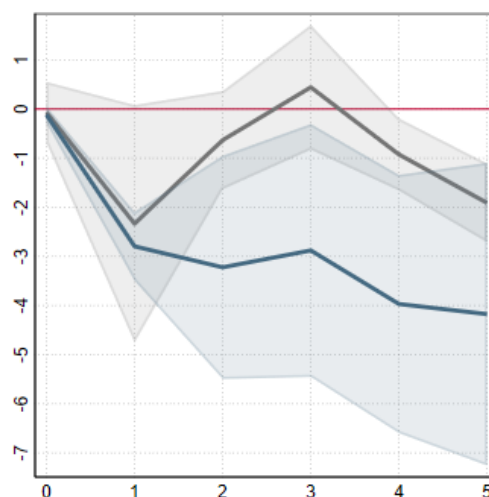


Figure B.27: Sub-sample analysis with 90^{th} - 10^{th} percentile difference



Notes: Upper LHS-panel shows the baseline and the specifications including the population-weighted standard deviation, 80^{th} - 20^{th} percentile difference, and 90^{th} - 10^{th} percentile difference in blue, green, and black, respectively. The other panels show the estimates from the sub-sample analysis for the bottom ten percent (in blue) and top ninety percent (in grey) of the per-capita GDP distribution for specifications including the respective regional dispersion measure as an additional control. For further detail, see notes to Figure B.23.

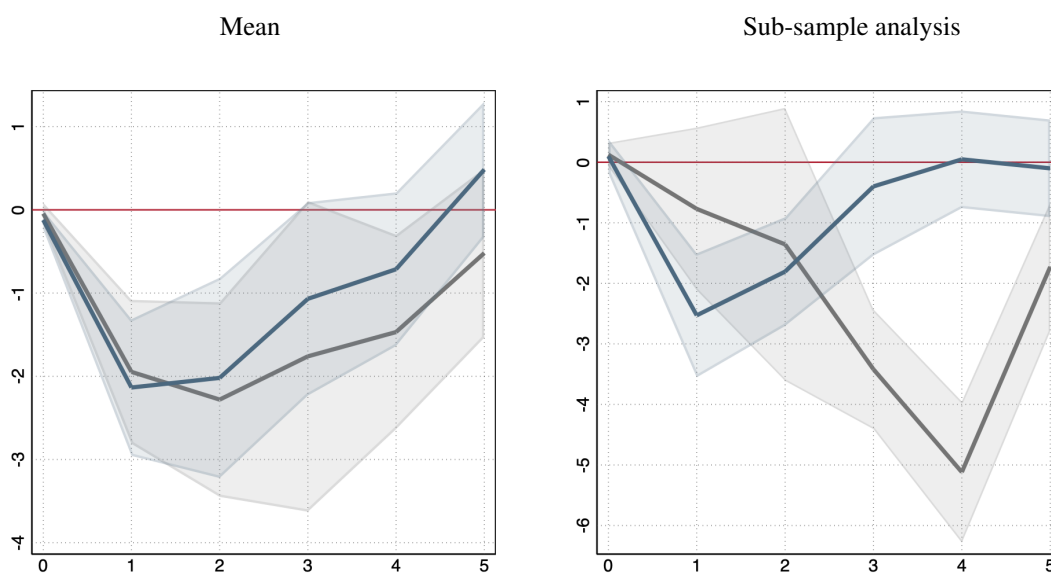
Tesar (2020) for a recent analysis documenting differential effects of exchange rate fluctuations across US regions).²³ Further, we include the long-term government bond yield spread vis-à-vis Germany in the list of explanatory variables to control for differences in country-specific

²³Similar effects appear plausible also for the euro area; for instance, Lane and Stracca (2018) document heterogeneity in exchange rate pass through across euro area countries and, given the large differences in economic structures at the NUTS3 level, such heterogeneity is likely to arise also here.

risk premia, which played a dominant role especially during the euro area sovereign debt crisis.

With some nuances, the main conclusions from our baseline estimations also carry over when including these additional regressors (Figure B.28). For the mean, the IRFs display very similar patterns up to horizon $h = 3$. After that point, the model including the additional variables yields a somewhat faster recovery of the output losses, but the difference to the baseline is again statistically insignificant. The sub-sample analysis, in turn, confirms the differential response across the GDP-distribution, characterized by a significantly stronger and more persistent contraction in economic activity in the lower than in the upper parts.

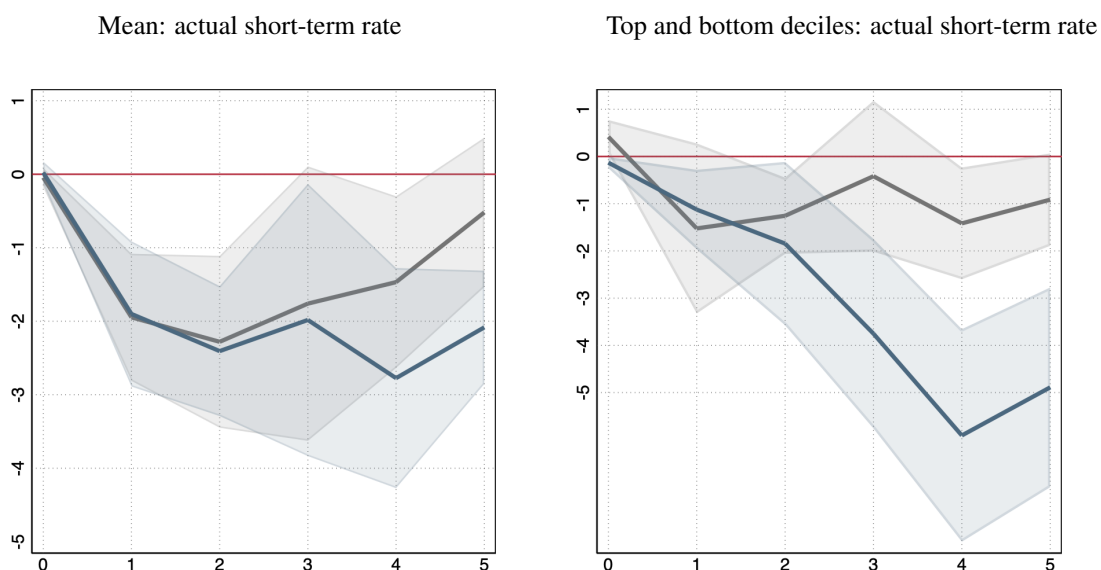
Figure B.28: Additional covariates



See notes to Figure B.23.

As a further specification check, we account for serial correlation in the residuals of the (shadow) short-term interest rates (Figure B.29). As these residuals serve to identify exogenous policy changes, such serial correlation may cast doubt on their being unanticipated. We therefore follow Ramey (2016) in extending the set of explanatory variables by the lagged residual of the policy indicator. The resultant IRFs point to a similar contraction at the mean as in the baseline, albeit again with a somewhat greater persistence, while confirming the pronounced heterogeneity across different parts of the distribution.

Figure B.29: Inclusion of lagged residuals



See notes to Figure B.23.

Appendix B.4. Alternative sample periods

The results shown so far have been estimated on our baseline sample running from 1999 through 2014. The sample starting date reflects our focus on euro area monetary policy, which has been assigned to the ECB since then. The endpoint is motivated by a structural break in monetary policy conduct that arose in 2015. Until then, the ECB had primarily steered its monetary policy stance via short-term interest rates. And while it had also resorted to other, ‘non-standard’, monetary policy measures since 2010, these had mainly served to preserve the smooth transmission of short-term interest rate changes, rather than acting as independent instruments to alter the stance. By contrast, the ECB’s introduction of large-scale asset purchases in January 2015 was explicitly motivated by the aim to achieve a further monetary policy easing, given the insufficient space to cut short-term interest rates, which were approaching their effective lower bound (ECB, 2015). The presence of multiple stance-related policy instruments, in turn, is difficult to capture in our baseline empirical setup.

We hence extend the sample period to include observations preceding the introduction of the euro and following the introduction of large-scale asset purchases. These extensions allow

Figure B.30: Long sample (blue) vs. pre-crisis sample (grey) vs. baseline sample (dashed blue)

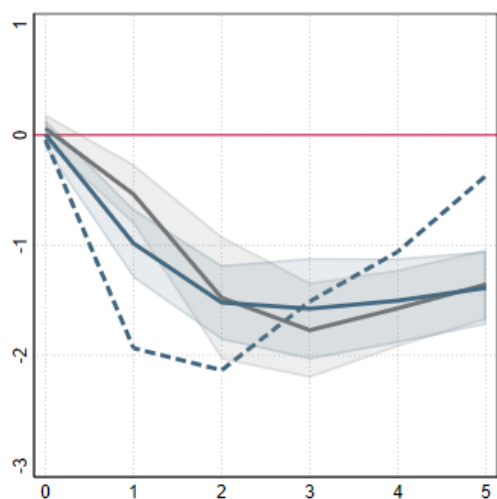
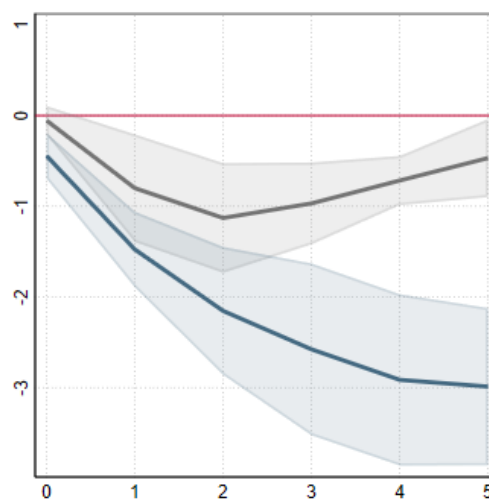


Figure B.31: Long sample: bottom vs. top deciles for GDP



Notes: Vertical axis refers to impact of a 100 basis point rate hike on the respective dependent variable (in %), horizontal axis refers to horizon of IRF (in years). In top-LHS, IRFs are for GDP during the extended period 1990-2014 (in blue), the pre-crisis period 1990-2007 (in grey) and the baseline period 1999-2014 (in dashed blue); in top-RHS, IRFs are for GDP during the extended period 1990-2014 at the bottom ten percent (in blue) and top ten percent (in grey) of the per-capita GDP distribution. Bottom-LHS shows the corresponding IRFs for employment and bottom-RHS for productivity, both over the extended period 1990-2014. Solid lines denote point estimates and shaded areas denote 90% confidence bands.

us to cater for the following considerations. First, the concept of hysteresis generally refers to adjustments in the economy over a medium- to longer-term horizon. As such, it seems important to also analyse the hysteresis patterns over a longer period, even though this requires us to impose some additional assumptions in the construction of the main variables for the period prior to the introduction of the euro (see below). Second, an earlier starting point raises the number of observations referring to the period prior to the Global Financial Crisis (GFC), which escalated in 2008. Since the GFC may have triggered structural shifts in key economic relationships, it is interesting to also estimate the model for the pre-crisis period. Third, it is worth exploring whether our findings also hold for broader measures of the monetary policy stance that capture the shift in policy conduct since 2015.

The first year for which the European Regional Database reports GDP at the NUTS3 level for a comprehensive set of regions is 1990, which we thus adopt as the alternative starting point for the sample. To ensure consistency with the previous research design, this sample extension requires us to back-cast some of the key variables in equation 1, including the policy-controlled

short-term interest rate i_t , with which we measure the euro area monetary policy stance. For the years prior to 1999, when national central banks were still able to set their own policy rates, we approximate this euro-area wide policy stance with the GDP-weighted average of the respective short-term interest rates at national level.²⁴ Likewise, we back-cast the euro area HICP with its GDP-weighted national average and euro area GDP with the sum of national GDPs over the pre-euro period. We then use this extended sample to reestimate the baseline model, first, over the full period from 1990 to 2014 and, second, over the pre-crisis period from 1990 to 2007.

For both sample definitions, we again find the monetary policy tightening to trigger marked and persistent contractions in regional output at the sample mean (Figure B.30). GDP hits its trough in the second and third year after the shock in the long and pre-crisis sample, respectively. Compared to the euro-area sample period, shown by the dashed line, the contraction is slightly shallower in the initial parts of the IRF horizon, but then proves more persistent.²⁵ Further, the estimates over the longer sample again clearly document differential effects of monetary policy across the distribution, with regions in the lower parts experiencing a more pronounced and long-lasting contraction in response to a tightening shock (Figure B.31).

Finally, we re-estimate our baseline model for the period from 1999 to 2018, as the last year for which NUTS3-level data are currently available, and replace the 3-month Euribor with

²⁴Specifically, we use the nominal short-term interest rates from the AMECO database of the European Commission. The short-term interest rate is predominantly the three-month interbank rate, though this might vary slightly between countries during given years. As an alternative to the GDP-weighted average of national short-term interest rates, we also experimented with a back-casting procedure that simply uses the German short-term rate as the measure of the euro area monetary policy stance over the period prior to euro introduction – a choice that may be motivated with the benchmark role of German Bundesbank for the policy stance of other European central banks in the run-up to euro introduction. But we found this alternative choice to yield almost identical results to that based on GDP-weighted averages.

²⁵While the point estimates remain statistically significant in $h = 5$, they converge back to zero when extending the IRF horizon for the long sample.

Figure B.32: Extended sample (1999-2018, blue) vs. baseline sample (1999-2014, grey)

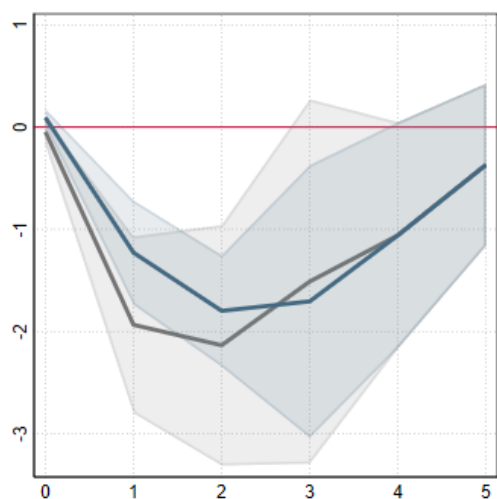
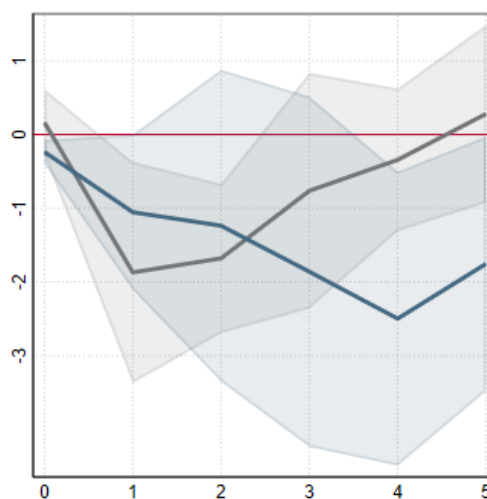


Figure B.33: Extended sample (1999-2018): bottom vs. top deciles for GDP



Notes: Vertical axis refers to impact of a 100 basis point rate hike on regional GDP (in %), horizontal axis refers to horizon of IRF (in years). On the LHS, grey line and shaded areas refer to the baseline sample period 1999-2014. Blue line and shaded areas refer to the sample period 1999-2018. Solid lines denote point estimates and shaded areas denote 90% confidence bands.

a ‘shadow’ short-term interest over the last three years of the sample. As such, we follow a recent literature that proposes shadow rates as a proxy of the monetary policy stance in the presence of lower bound constraints and non-standard measures (Krippner, 2015; Wu and Xia, 2017; Lemke and Vladu, 2017).²⁶

At the mean, the estimates over the extended sample period exhibit very similar patterns as in the baseline and the confidence intervals overlap throughout the IRF horizon (Figure B.32). Moreover, the extended sample estimates again point to a heterogeneous footprint of monetary policy. Whereas richer regions recover from the monetary-policy induced GDP contraction after a few years, the impact on poorer regions proves persistent and still remains statistically significant by the end of the horizon (Figure B.33).

²⁶In terms of implementation, we rely on the shadow rate developed by Lemke and Vladu (2017) and add the cumulative changes in that rate to the 2014 level of the short-term interest rate deployed in our baseline estimations.

Appendix C. Analytical results

Appendix C.1. Proof of proposition 1

Steady state output (in log-levels) is given by

$$y = \left(\frac{\phi}{1-\rho} + 1 - \alpha \right) n. \quad (\text{C.1})$$

Hence

$$\frac{\partial y}{\partial \rho} = \frac{\phi}{(1-\rho)^2} n + \left(\frac{\phi}{1-\rho} + 1 - \alpha \right) \frac{\partial n}{\partial \rho}. \quad (\text{C.2})$$

The first term is negative since $0 < N < 1$, which means $n < 0$. Thus, $\frac{\partial y}{\partial \rho}$ is unambiguously negative if $\frac{\partial n}{\partial \rho} < 0$.

Steady state employment is given by

$$\xi (1-N)^{-\varphi} = (1-\alpha) N^{\left(\frac{\phi}{1-\rho} + 1 - \alpha \right) (1-\sigma) - 1}. \quad (\text{C.3})$$

Taking logs

$$\log \xi - \varphi \log (1-N) = \log (1-\alpha) + \left(\left(\frac{\phi}{1-\rho} + 1 - \alpha \right) (1-\sigma) - 1 \right) \log N. \quad (\text{C.4})$$

Total differentiation with respect to ρ yields

$$\varphi \frac{1}{1-N} \frac{\partial N}{\partial \rho} = \frac{\phi}{(1-\rho)^2} (1-\sigma) \log N + \left(\left(\frac{\phi}{1-\rho} + 1 - \alpha \right) (1-\sigma) - 1 \right) \frac{1}{N} \frac{\partial N}{\partial \rho}. \quad (\text{C.5})$$

Rearranging gives

$$\frac{\partial N}{\partial \rho} = \frac{\frac{\phi}{(1-\rho)^2} (1-\sigma) \log N}{\varphi \frac{1}{1-N} - \left(\left(\frac{\phi}{1-\rho} + 1 - \alpha \right) (1-\sigma) - 1 \right) \frac{1}{N}}. \quad (\text{C.6})$$

Since we imposed the condition $\alpha > \phi/(1-\rho)$ to ensure non-increasing returns to scale, the term $\left(\left(\frac{\phi}{1-\rho} + 1 - \alpha \right) (1-\sigma) - 1 \right)$ is negative and the denominator is positive. Thus, since $0 < N < 1$ (i.e., $\log N < 0$), we have that $\frac{\partial N}{\partial \rho} < 0$ so long as $\sigma < 1$. This completes the proof.

Appendix C.2. Proof of proposition 2

The labour market clearing and TFP equations in log-deviations from steady state are given by

$$0 = (1-\sigma) \hat{a}_t - \mathcal{A} \hat{n}_t, \quad (\text{C.7})$$

$$\hat{a}_t = \rho \hat{a}_{t-1} + \phi \hat{n}_t, \quad (\text{C.8})$$

where $\hat{x}_t \equiv x_t - x$ denotes log-deviations from steady state, $\mathcal{A} \equiv 1 + \tilde{\varphi} - (1-\alpha)(1-\sigma) > 0$ as long as $\sigma < 1$, and where $\tilde{\varphi} \equiv \frac{N}{1-N} \varphi$.

Correlation: Since $\hat{n}_t = \frac{1-\sigma}{\mathcal{A}} \hat{a}_t$, it follows that $\text{corr}(n_t, a_t) > 0$ as long as $\sigma < 1$.

Persistence: First, suppose that employment, \hat{n}_t , is perturbed by a non-TFP shock, u_t , such that $\hat{n}_t = \mathcal{B} \hat{a}_t + u_t$, where u_t follows an AR(1) process $u_{t+1} = \theta u_t + \varepsilon_{t+1}$, where ε_{t+1} is an iidN innovation and $\mathcal{B} \equiv \frac{1-\sigma}{\mathcal{A}}$. Next, we can rewrite the equation for \hat{a}_t and the AR(1) process using lag operators as $\hat{a}_t = \frac{\phi}{1-\rho L} \hat{n}_t$ and $u_t = \frac{1}{1-\theta L} \varepsilon_t$, respectively. Rewriting in terms of n_t , we

get

$$n_t = \frac{\mathcal{B}\phi}{1 - \rho L} n_t + \frac{1}{1 - \theta L} e_t, \quad (\text{C.9})$$

$$(1 - \rho L)(1 - \theta L)n_t = (1 - \theta L)\mathcal{B}\phi n_t + (1 - \rho L)e_t, \quad (\text{C.10})$$

$$n_t = \frac{\rho + \theta - \theta\mathcal{B}\phi}{1 - \mathcal{B}\phi} n_{t-1} - \frac{\rho\theta}{1 - \mathcal{B}\phi} n_{t-2} + \tilde{e}_t - \rho\tilde{e}_{t-1}, \quad (\text{C.11})$$

$$= s_1 n_{t-1} + s_2 n_{t-2} + \tilde{e}_t + c\tilde{e}_{t-1}, \quad (\text{C.12})$$

where $s_1 \equiv \frac{\rho + \theta - \theta\mathcal{B}\phi}{1 - \mathcal{B}\phi}$, $s_2 \equiv -\frac{\rho\theta}{1 - \mathcal{B}\phi}$, and $c \equiv -\rho$. Thus, n_t follows an *ARMA*(2,1) process.²⁷

Using standard results, it is straightforward to calculate the first-order autocorrelation of an *ARMA*(2,1) process:

$$\text{ac}(\cdot) = \frac{(1 + c(s_1 + c))s_1 + (1 - s_2^2)c}{s_1(1 + s_2)c + (1 + c(s_1 + c))(1 - s_2)}. \quad (\text{C.13})$$

When $\rho = 0$, the autocorrelation of employments follows the exogenous process: $\text{ac}(n_t) = \theta$.

Next we calculate the partial derivative with respect to ρ near $\rho = 0$:

$$\left. \frac{\partial \text{ac}(n_t)}{\partial \rho} \right|_{\rho=0} = \mathcal{B} \frac{(1 - \theta^2)\phi}{1 - \mathcal{B}\phi}. \quad (\text{C.14})$$

Thus, $\left. \frac{\partial \text{ac}(n_t)}{\partial \rho} \right|_{\rho=0} > 0$ (i.e., employment becomes more persistent when ρ increases) if $\mathcal{B} > 0$, which requires $\sigma < 1$.

Using the equation for output, $y_t = a_t + (1 - \alpha)n_t$, and labour productivity, $lp_t = y_t - n_t =$

²⁷Since the autocorrelation function is independent of the variance of ε_t , we have renormalized this variable and denoted it as \tilde{e}_t to avoid unnecessary terms.

$a_t - \alpha n_t$, we can also write y_t and lp_t as $ARMA(2, 1)$ processes. Specifically

$$y_t = \frac{\theta + \rho - \theta \mathcal{B}\phi}{1 - \mathcal{B}\phi} y_{t-1} - \frac{\rho\theta}{1 - \mathcal{B}\phi} y_{t-2} + \tilde{\varepsilon}_t - \frac{\rho}{1 + \frac{\phi}{1-\alpha}} \tilde{\varepsilon}_{t-1}, \quad (\text{C.15})$$

$$lp_t = \frac{\theta + \rho - \theta \mathcal{B}\phi}{1 - \mathcal{B}\phi} lp_{t-1} - \frac{\rho\theta}{1 - \mathcal{B}\phi} lp_{t-2} + \varepsilon_t - \frac{\rho}{1 + \frac{\phi}{-\alpha}} \varepsilon_{t-1}. \quad (\text{C.16})$$

In both cases, $ac(y_t) = ac(lp_t) = \theta$ when $\rho = 0$. The partial derivatives with respect to ρ near $\rho = 0$ are given by

$$\left. \frac{\partial ac(y_t)}{\partial \rho} \right|_{\rho=0} = \frac{1 + \mathcal{B}(1 - \alpha)}{1 - \alpha + \phi} \frac{(1 - \theta^2)\phi}{1 - \mathcal{B}\phi}, \quad (\text{C.17})$$

$$\left. \frac{\partial ac(lp_t)}{\partial \rho} \right|_{\rho=0} = \frac{1 - \mathcal{B}\alpha}{-\alpha + \phi} \frac{(1 - \theta^2)\phi}{1 - \mathcal{B}\phi}, \quad (\text{C.18})$$

$$(\text{C.19})$$

For output, the key term is $1 + \mathcal{B}(1 - \alpha)$. The requirement for $\left. \frac{\partial ac(y_t)}{\partial \rho} \right|_{\rho=0} > 0$ is not as strong as for employment. Only if \mathcal{B} is sufficiently negative (caused by an extremely large σ) would the counterbalancing force of increased labour supply in response to falling TFP cause output hysteresis to fall as ρ increases. For labour productivity, the key term is $\frac{1 - \mathcal{B}\alpha}{-\alpha + \phi}$.

Appendix C.3. Real Business Cycle (RBC) example

This example is similar to the one in the main text, with the key difference that we employ a real business cycle (RBC) and generate hysteresis by relating TFP to the level of region-wide investment spending.

Time, t , is discrete and a region i is populated by an infinitely lived representative household with preferences over consumption, c_t , given by $\mathbb{E}_t \sum_{t=0}^{\infty} \beta^t \log(c_t)$, where $\beta \in (0, 1)$ is the discount factor. The household resource constraint is given by $c_t + i_t = y_t = A_t k_{t-1}^\alpha$, where A_t is total factor productivity (TFP), i_t is investment, k_t is the capital stock, and $\alpha \in (0, 1)$. Capital

evolves according to $k_t = (1 - \delta)k_{t-1} + i_t$ although we restrict attention to the case of full capital depreciation in every period (in which case $\delta = 1$). TFP is assumed to be endogenous and increasing in aggregate (region-level) investment spending, I_t , which individual households take as given. Concretely, we assume the evolution of TFP is given by $A_t = A_{t-1}^{\rho_i} I_{t-1}^\phi$, where $\rho_i, \phi \in (0, 1)$. Note that ρ_i is again the only region-specific structural parameter in the set-up of the model. All other parameters are common across regions. Finally, we impose the parameter restriction $\frac{\phi}{1-\rho_i} + \alpha < 1$ which ensures stationarity.

The first-order condition of the household is given by $1 = \beta \mathbb{E}_t (c_t/c_{t+1}) A_{t+1} \alpha k_t^{\alpha-1}$. In equilibrium, all household variables correspond to their aggregate values. Hence $x_t = X_t$ for $x_t = \{c_t, i_t, y_t, k_t\}$. Furthermore, it is straightforward to show via guess and verify that, in equilibrium, investment and consumption are fixed fractions ($\alpha\beta$ and $1 - \alpha\beta$, respectively) of output. Thus, the evolution of the economy is fully described by the joint dynamics of the capital stock, $\hat{k}_t = \log(\alpha\beta) + \rho_i \hat{a}_{t-1} + (\phi + \alpha) \hat{k}_{t-1}$, and TFP, $\hat{a}_t = \rho_i \hat{a}_{t-1} + \phi \hat{k}_{t-1}$, where $\hat{x} = \log(X)$. In log-deviations from steady state, $\tilde{k}_t = \rho_i \tilde{a}_{t-1} + (\phi + \alpha) \tilde{k}_{t-1}$ and $\tilde{a}_t = \rho_i \tilde{a}_{t-1} + \phi \tilde{k}_{t-1}$, where $\tilde{x}_t = \hat{x}_t - \hat{x}$.

We are now ready to present our two key propositions. The first concerns the steady state of the model and the second concerns the persistence of the model dynamics when transitioning back to steady state.

Proposition 3 *Steady state output is decreasing in ρ_i*

Proposition 4 *The unconditional autocorrelation of output is increasing in ρ_i*

The two propositions, taken together, are consistent with the main finding of this paper that deviations of output are more persistent in Low income regions. Note that our model and the propositions that follow are very general. We do not need to specify the shock that drives the economy away from steady state.

Proof of Proposition 3. The steady state of TFP is given by $\hat{a} = \frac{\phi}{1-\rho_i} \hat{k}$. Substituting into the capital accumulation equation and rearranging gives $\hat{k} = \frac{1}{1-\alpha-\frac{\phi}{1-\rho_i}} \log(\alpha\beta)$. Substituting into the production function gives steady state output as $\hat{y} = \frac{\alpha(1-\rho_i)+\phi}{(1-\alpha)(1-\rho_i)-\phi} \log(\alpha\beta)$. It is straightforward to show that $\partial\hat{k}/\partial\rho_i, \partial\hat{y}/\partial\rho_i < 0$. ■

Proof of Proposition 4. First, rewrite the TFP evolution equation with lag operators as follows: $\tilde{a}_t = \frac{\phi}{(1-\rho_i L)} \tilde{k}_{t-1}$. Next, substitute this expression into the capital accumulation equation and expand so that it becomes an $AR(2)$ in \tilde{k}_t only, as follows:

$$\begin{aligned}\tilde{k}_t &= \frac{\phi}{(1-\rho_i L)} \tilde{k}_{t-1} + \alpha \tilde{k}_{t-1}, \\ (1-\rho_i L) \tilde{k}_t &= \phi \tilde{k}_{t-1} + (1-\rho_i L) \alpha \tilde{k}_{t-1}, \\ \tilde{k}_t &= (\rho_i + \phi + \alpha) \tilde{k}_{t-1} - \rho_i \alpha \tilde{k}_{t-2}.\end{aligned}$$

A standard result of an $AR(2)$ process of the form $x_t = \theta_1 x_{t-1} + \theta_2 x_{t-2}$ is a first-order autocorrelation given by $ac(x_t) \equiv \frac{\mathbb{E}(x_t x_{t-1})}{\mathbb{E}(x_t^2)} = \frac{\theta_1}{1-\theta_2}$. Hence, $ac(\tilde{k}_t) = \frac{\rho_i + \phi + \alpha}{1 + \rho_i \alpha}$. It is straightforward to show that $\partial ac(\tilde{k}_t) / \partial \rho_i > 0$. Since $\tilde{y}_t = \tilde{k}_t$, it follows that $\partial ac(\tilde{y}_t) / \partial \rho_i > 0$. ■

Appendix D. A currency union of small-open economy regions

The model follows closely [Gali and Monacelli \(2008\)](#). The currency union consists of a continuum of small open economies (“regions”) of unit mass. We assume the regions share the same structural parameters, except one—the persistence of endogenous total factor productivity changes. Furthermore, we assume the law-of-one-price holds across the union and that households can trade a complete set of state-contingent claims.

Households The representative household in region i has preferences given by

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \frac{(C_t^i)^{1-\sigma}}{1-\sigma} + \xi \frac{(1-N_t^i)^{1-\varphi}}{1-\varphi}, \quad (\text{D.1})$$

where N_t^i denotes hours worked and C_t^i is a composite consumption index, constructed of the following objects:

$$C_t^i \equiv \frac{(C_{i,t}^i)^{1-\omega} (C_{F,t}^i)^{\omega}}{(1-\omega)^{1-\omega} \omega^{\omega}}, \quad C_{F,t}^i \equiv \exp \int_0^1 \log(C_{f,t}^i) df, \quad C_{f,t}^i \equiv \left(\int_0^1 C_{f,t}^i(j)^{\frac{\varepsilon-1}{\varepsilon}} dj \right)^{\frac{\varepsilon}{\varepsilon-1}}, \quad (\text{D.2})$$

where $C_{F,t}^i$ is a consumption index of imported goods and $C_{f,t}^i$ is an index of goods produced by region f and consumed by region i , where $j \in [0, 1]$ denotes the type of good. The parameter ω is an index of openness (where decreasing ω reflects greater home bias), and $\varepsilon > 1$ is the elasticity of substitution between varieties in a given region.

The household maximizes (D.1) subject to the budget constraint given by

$$\int_0^1 P_t^i(j) C_{i,t}^i(j) dj + \int_0^1 \int_0^1 P_t^f(j) C_{f,t}^i(j) dj df + \mathbb{E}_t(Q_{t+1} D_{t+1}^i) = D_t^i + W_t^i N_t^i - T_t^i, \quad (\text{D.3})$$

where $P_t^f(j)$ is the price of good j produced by region f ; D_{t+1}^i denotes a random variable

indicating the number of assets purchased in period t to be delivered in each state of period $t + 1$ and Q_{t+1} is the period t price of an asset that pays one unit in a particular state in $t + 1$ divided by the time- t conditional probability of that state occurring. W_t^i are wages and T_t^i are lump-sum taxes and dividends (in nominal terms).

The optimal allocation of expenditure yields demand curves given by

$$\frac{C_{f,t}^i(j)}{C_{f,t}^i} = \left(\frac{P_t^f(j)}{P_t^f} \right)^{-\varepsilon}, \quad \frac{C_{f,t}^i}{C_{F,t}^i} = \left(\frac{P_t^f}{P_t^*} \right)^{-1}, \quad (\text{D.4})$$

where $P_t^f \equiv \left(\int_0^1 P_t^f(j)^{1-\varepsilon} dj \right)^{\frac{1}{1-\varepsilon}}$ is both country f 's domestic price index and the price index for the goods imported from region f by country i , and $P_t^* \equiv \exp \int_0^1 \log(P_t^f) df$ is the union-wide price index. It follows that $P_t^f C_{f,t}^i = \int_0^1 P_t^f(j) C_{f,t}^i(j) dj$ and $P_t^* C_{F,t}^i = \int_0^1 P_t^f C_{f,t}^i df$. The optimal allocation of expenditures between domestic and imported good is given by

$$\frac{C_{i,t}^i}{C_t^i} = (1 - \omega) \left(\frac{P_t^i}{P_{c,t}^i} \right)^{-1} \quad \text{and} \quad \frac{C_{F,t}^i}{C_t^i} = \omega \left(\frac{P_t^*}{P_{c,t}^i} \right)^{-1}, \quad (\text{D.5})$$

where $P_{c,t}^i \equiv (P_t^i)^{1-\omega} (P_t^*)^\omega$ denotes the Consumer Price Index (CPI) for region i . Total consumption expenditure in region i is given by $P_{c,t}^i C_t^i = P_t^i C_{i,t}^i + P_t^* C_{F,t}^i$. Thus, the period budget constraint can be rewritten as

$$P_{c,t}^i C_t^i + \mathbb{E}_t(Q_{t+1} D_{t+1}^i) = D_t^i + W_t^i N_t^i - T_t^i. \quad (\text{D.6})$$

The respective optimality conditions for households' labour/leisure and consumption/savings

decisions in region i are given by

$$\xi (C_t^i)^\sigma (1 - N_t^i)^{-\varphi} = \frac{W_t^i}{P_{c,t}^i}, \quad (\text{D.7})$$

$$\beta \left(\frac{C_{t+1}^i}{C_t^i} \right)^{-\sigma} \frac{1}{\Pi_{c,t+1}^i} = Q_{t+1}, \quad (\text{D.8})$$

where $\Pi_{c,t}^* = P_{c,t}^i / P_{c,t-1}^i$. Taking expectations of (D.8) gives the conventional Euler equation

$$\beta R_t^* \mathbb{E}_t \left(\frac{C_{t+1}^i}{C_t^i} \right)^{-\sigma} \frac{1}{\Pi_{c,t+1}^i} = 1, \quad (\text{D.9})$$

where $R_t^* = 1 / \mathbb{E}_t Q_{t+1}$.

Next, define the bilateral terms of trade between region i and f as $S_{f,t}^i \equiv P_t^f / P_t^i$ and the effective terms of trade for region i as $S_t^i = P_t^* / P_t^i = \exp \int_0^1 \log \left(S_{f,t}^i \right) df$. The CPI and domestic price levels are related by $P_{c,t}^i = P_t^i (S_t^i)^\omega$ and hence $\Pi_{c,t}^i = \Pi_t^i (\Delta S_t^i)^\omega$. For the currency union as a whole $P_{c,t}^* = P_t^*$ and hence $\Pi_{c,t}^* = \Pi_t^*$.

International risk sharing Under the assumption of complete markets for state-contingent securities, an analogous first-order condition to (D.8) for region f is given by

$$\beta \left(\frac{C_{t+1}^f}{C_t^f} \right)^{-\sigma} \frac{1}{\Pi_{c,t}^f} = Q_{t,t+1}. \quad (\text{D.10})$$

Combining first-order conditions gives

$$\frac{C_t^i}{C_t^f} = \frac{C_{t+1}^i}{C_{t+1}^f} \frac{P_{c,t+1}^i}{P_{c,t+1}^f} \frac{P_{c,t}^i}{P_{c,t}^f}. \quad (\text{D.11})$$

This has the form $x_t = x_{t+1} y_{t+1} / y_t$. Rearranging gives $x_t = x_{t-1} y_{t-1} / y_t$. Iterating backwards

yields $x_t = x_0 y_0 / y_t$. Hence, we can rewrite (D.11) as

$$\frac{C_t^i}{C_t^f} = v_{i,f} \frac{P_{c,t}^f}{P_{c,t}^i} = v_{i,f} (S_{f,t}^i)^{1-\omega}. \quad (\text{D.12})$$

By assuming symmetric initial conditions, $v_{i,f} = v = 1$. Integrating over f , we obtain

$$C_t^i = C_t^* (S_t^i)^{1-\omega}, \quad (\text{D.13})$$

where $C_t^* \equiv \exp \int_0^1 \log C_t^f df$.

Firms Each region has a unit continuum of firms, each producing a differentiated good with $Y_t^i(j) = A_t^i (N_t^i(j))^{1-\alpha}$ where A_t^i is region-specific. Real marginal cost is given by

$$MC_t^i(j) = \left(\frac{1-\tau^i}{1-\alpha} \right) \frac{W_t^i}{P_t^i} \frac{1}{A_t^i (N_t^i(j))^{-\alpha}}. \quad (\text{D.14})$$

where τ^i is a constant employment subsidy.

The aggregate output index for region i is given by

$$Y_t^i \equiv \left[\int_0^1 Y_t^i(j)^{\frac{\varepsilon-1}{\varepsilon}} dj \right]^{\frac{\varepsilon}{\varepsilon-1}}, \quad (\text{D.15})$$

and the amount of labour hired is given by

$$N_t^i \equiv \int_0^1 N_t^i(j) dj = \left(\frac{Y_t^i}{A_t^i} \right)^{\frac{1}{1-\alpha}} \int_0^1 \left(\frac{P_t^i(j)}{P_t^i} \right)^{-\frac{\varepsilon}{1-\alpha}} dj. \quad (\text{D.16})$$

Price setting Firms set prices as in Calvo (1983), where $1 - \theta$ is the probability that a firms can set a new price. The fact that each individual region is open does not affect the form of the equation relating domestic inflation to real marginal cost. As such, the dynamics of domestic

inflation in terms of real marginal costs in region i is given by

$$\hat{\pi}_t^i = \beta \mathbb{E}_t \hat{\pi}_{t+1}^i + \lambda \hat{m}c_t^i, \quad (\text{D.17})$$

where $\hat{m}c_t^i = mc_t^i - \mu$ denotes the log-deviation of real marginal costs from its steady state, and

$$\lambda = \frac{(1-\beta\theta)(1-\theta)}{\theta} \Theta, \text{ where } \Theta \equiv \frac{1-\alpha}{1-\alpha+\alpha\varepsilon}.$$

Aggregate dynamics The clearing of the market for good j in region i is given by

$$\begin{aligned} Y_t^i(j) &= C_{i,t}^i(j) + \int_0^1 C_{f,t}^i(j) df \\ &= \left(\frac{P_t^i(j)}{P_t^i} \right)^{-\varepsilon} \left(C_{i,t}^i + \int_0^1 C_{i,t}^f df \right), \\ &= \left(\frac{P_t^i(j)}{P_t^i} \right)^{-\varepsilon} \left((1-\omega) \left(\frac{P_t^i}{P_{c,t}^i} \right)^{-1} C_t^i + \omega \int_0^1 \left(\frac{P_t^i}{P_{c,t}^f} \right)^{-1} C_t^f df \right), \\ &= \left(\frac{P_t^i(j)}{P_t^i} \right)^{-\varepsilon} \left((1-\omega) (S_t^i)^\omega C_t^i + \omega \int_0^1 (S_t^i)^\omega (S_{f,t}^i)^{1-\omega} C_t^f df \right), \\ &= \left(\frac{P_t^i(j)}{P_t^i} \right)^{-\varepsilon} C_t^i (S_t^i)^\omega, \end{aligned} \quad (\text{D.18})$$

and hence the market clearing condition is $Y_t^i = C_t^i (S_t^i)^\omega$.

Endogenous total factor productivity Finally, we arrive at the key feature that distinguishes regions. Denote $U_t^i = 1 - N_t^i$ as the unemployment rate. We assume that total factor productivity (TFP) accumulation follows $A_t^i = (A_{t-1}^i)^{\rho^i} (1 - U_t^i)^\phi$, where A_t^i is increasing in past TFP and decreasing in the unemployment rate. In other words, unemployment causes a de-accumulation of skills or human capital.

Equilibrium dynamics Next, we rewrite the model in log-deviations, where $x \equiv \log(X)$ and

$\hat{x}_t \equiv x_t - x$. The labour supply condition and Euler equation are then

$$\sigma \hat{c}_t^i + \tilde{\varphi} \hat{n}_t^i = \hat{w}_t^i - \hat{p}_{c,t}^i, \quad (\text{D.19})$$

$$\hat{c}_t^i = \mathbb{E}_t \hat{c}_{t+1}^i - \frac{1}{\sigma} (\hat{r}_t^* - \mathbb{E}_t \pi_{t+1}^i), \quad (\text{D.20})$$

where $\tilde{\varphi} \equiv \varphi \frac{N^i}{1-N^i}$. The CPI and domestic price levels and inflation rates are related by

$$\hat{p}_{c,t}^i = \hat{p}_t^i + \omega \hat{s}_t^i, \quad \hat{\pi}_{c,t}^i = \hat{\pi}_t^i + \omega \Delta \hat{s}_t^i. \quad (\text{D.21})$$

International risk sharing ensures

$$\hat{c}_t^i = \hat{c}_t^* + (1 - \omega) \hat{s}_t^i. \quad (\text{D.22})$$

Equilibrium variations in $\int_0^1 f(P_t^i(j)/P_t^i) dj$ are of second-order. Thus, up to a first-order approximation, the following relationships hold:

$$\hat{m}c_t^i = \hat{w}_t^i - \hat{p}_t^i - \hat{a}_t^i + \alpha \hat{n}_t^i, \quad (\text{D.23})$$

$$\hat{y}_t^i = \hat{a}_t^i + (1 - \alpha) \hat{n}_t^i. \quad (\text{D.24})$$

The market clearing condition is

$$\hat{y}_t^i = \hat{c}_t^i + \omega \hat{s}_t^i = \hat{c}_t^* - (\hat{p}_t^i - \hat{p}_t^*). \quad (\text{D.25})$$

Integrating over (D.25) yields

$$\hat{y}_t^* = \hat{c}_t^*. \quad (\text{D.26})$$

Thus, the union-wide IS curve is given by

$$\hat{y}_t^* = \mathbb{E}_t \hat{y}_{t+1}^* - \frac{1}{\sigma} (\hat{r}_t^* - \mathbb{E}_t \pi_{t+1}^*). \quad (\text{D.27})$$

Marginal cost can be rewritten as follows:

$$\begin{aligned} \hat{m}c_t^i &= \hat{w}_t^i - \hat{p}_t^i - \hat{a}_t^i + \alpha \hat{n}_t^i, \\ &= (\hat{w}_t^i - \hat{p}_{c,t}^i) + (\hat{p}_{c,t}^i - \hat{p}_t^i) - \hat{a}_t^i + \alpha \hat{n}_t^i, \\ &= \sigma \hat{c}_t^i + \tilde{\varphi} \hat{n}_t^i + \omega s_t^i - \hat{a}_t^i + \alpha \hat{n}_t^i, \\ &= \sigma (\hat{y}_t^i - \omega s_t^i) + (\tilde{\varphi} + \alpha) \hat{n}_t^i + \omega s_t^i - \hat{a}_t^i, \\ &= \sigma (\hat{y}_t^i - \omega s_t^i) + \left(\frac{\tilde{\varphi} + \alpha}{1 - \alpha} \right) (\hat{y}_t^i - \hat{a}_t^i) + \omega s_t^i - \hat{a}_t^i, \\ &= \left(\sigma + \frac{\tilde{\varphi} + \alpha}{1 - \alpha} \right) \hat{y}_t^i - \left(\frac{1 + \tilde{\varphi}}{1 - \alpha} \right) \hat{a}_t^i + (1 - \sigma) \omega s_t^i. \end{aligned} \quad (\text{D.28})$$

Hence, the Phillips curve is given by

$$\hat{\pi}_t^i = \beta \mathbb{E}_t \hat{\pi}_{t+1}^i + \kappa_y \hat{y}_t^i - \kappa_a \hat{a}_t^i + \kappa_s \hat{s}_t^i, \quad (\text{D.29})$$

where $\kappa_y \equiv \lambda \left(\sigma + \frac{\tilde{\varphi} + \alpha}{1 - \alpha} \right)$, $\kappa_a \equiv \lambda \left(\frac{1 + \tilde{\varphi}}{1 - \alpha} \right)$ and $\kappa_s \equiv \lambda (1 - \sigma) \omega$.

Monetary policy Monetary policy is described by a simple Taylor-type rule

$$\hat{r}_t^* = \phi \hat{\pi}_t^* + u_t, \quad (\text{D.30})$$

where $u_{t+1} = \rho u_t + \sigma_u \varepsilon_{t+1}$ and ε_{t+1} is drawn from an *iid* standard normal distribution.

Equilibrium system Suppose there are two types of region: High (H) and Low (L) income regions, with mass w and $1 - w$, respectively. The full set of 13 endogenous variables are:

$\{\hat{\pi}_t^i, \hat{y}_t^i, \hat{a}_t^i, \hat{n}_t^i, \hat{s}_t^i\}$ and $\{\hat{r}_t^*, \hat{\pi}_t^*, \hat{y}_t^*\}$. The full system of 13 equations is given by:

1. Regional equations for $i = H, L$:

$$\pi_t^i = \beta \mathbb{E}_t \pi_{t+1}^i + \kappa_y \hat{y}_t^i - \kappa_d \hat{a}_t^i + \kappa_s \hat{s}_t^i, \quad (\text{D.31})$$

$$\hat{y}_t^i = \hat{y}_t^* + \hat{s}_t^i, \quad (\text{D.32})$$

$$\hat{y}_t^i = \hat{a}_t^i + (1 - \alpha) \hat{n}_t^i, \quad (\text{D.33})$$

$$\hat{a}_t^i = \rho^i \hat{a}_{t-1}^i + \phi \hat{n}_t^i. \quad (\text{D.34})$$

2. Definition

$$\Delta \hat{s}_t^H = \hat{\pi}_t^* - \hat{\pi}_t^H. \quad (\text{D.35})$$

3. Aggregation equations

$$\hat{\pi}_t^* = w \hat{\pi}_t^H + (1 - w) \hat{\pi}_t^L, \quad (\text{D.36})$$

$$\hat{y}_t^* = w \hat{y}_t^H + (1 - w) \hat{y}_t^L. \quad (\text{D.37})$$

4. Union wide equation

$$\hat{y}_t^* = \mathbb{E}_t \hat{y}_{t+1}^* - \frac{1}{\sigma} (\hat{r}_t^* - \mathbb{E}_t \pi_{t+1}^*). \quad (\text{D.38})$$

$$\hat{r}_t^* = \phi \hat{\pi}_t^* + u_t. \quad (\text{D.39})$$

Steady state We impose zero-inflation steady state $\Pi_c^i = \Pi = 1$ and set the employment

subsidy, τ^i to ensure the steady state is not distorted.²⁸ Then:

$$A^i = (N^i)^{\frac{\phi}{1-\rho^i}}, \quad (\text{D.40})$$

$$\xi (C^i)^\sigma (1 - N^i)^{-\varphi} = (1 - \alpha) A^i (N^i)^{-\alpha}, \quad (\text{D.41})$$

$$Y^i = C^i, \quad (\text{D.42})$$

$$Y^i = A^i (N^i)^{1-\alpha}. \quad (\text{D.43})$$

Combining these steady-state conditions, we obtain a non-linear expression in N^i :

$$(N^i)^{1 + \left(\frac{\phi}{1-\rho^i} - \alpha\right)(\sigma-1)} (1 - N^i)^{-\varphi} = \frac{1 - \alpha}{\xi}. \quad (\text{D.44})$$

Calibration The calibration is presented in Table D.4. The subjective discount factor, $\beta = 0.99$, implies a steady state annual real interest rate of 4%. The intertemporal elasticity of substitution (IES), $1/\sigma = 2$. An IES greater than 1 ensures that the wealth effect of changes in TFP does not dominate the substitution effect on labour supply of a change in the real wage. Hence, the effect of endogenously falling TFP is exacerbated by an additional fall in employment. For the High income region, we calibrate the model such that the steady state unemployment rate is 5% ($1 - N^H = 0.05$) and the Frisch elasticity of labour supply, $\tilde{\varphi}^H = 1$. This fixes the preference parameters φ and ξ . The labour share of income, $1 - \alpha = 0.8$. This parameter is set to less than one, despite production being absent capital or other variable inputs, because this ensures that labour productivity is time-varying as in the data. The elasticity of substitution across goods (in a given region), $\varepsilon = 3$, implies a steady state mark up of 1.5. The Calvo parameter, $\theta = 2/3$, implies that firms adjust prices on average every 3 quarters. House-

²⁸Since we do not study optimal monetary policy, this subsidy has no impact on the equilibrium dynamics.

holds are assumed to have a home bias in consumption, with $\omega = 0.5$. We assume that there are only two types of region (High and Low income regions), with an equal fraction of both ($w = 0.5$). Moreover, we assume one region type experiences hysteresis ($\rho^L = 0.925$) whereas the other does not ($\rho^H = 0$). Note that while we label these regions High and Low income, these designations are an endogenous outcome of the model. The unemployment elasticity of TFP, $\phi = 0.1$, implies that a 1 percentage point rise in the unemployment rate decreases TFP by 10 basis points.

Table D.4: Calibration

Parameter	Description	Value
β	Subjective discount factor	0.99
$1/\sigma$	Intertemporal elasticity of substitution	2
φ	Utility curvature on leisure	0.053
ξ	Utility weight on leisure	0.721
$1 - \alpha$	Labour share of income	0.8
ε	Elasticity of substitution	3
θ	Calvo parameter	0.667
ω	Home bias	0.5
w	Share of High income regions	0.5
ϕ	Unemployment elasticity of TFP	0.1
ρ^H	TFP persistence (High income region)	0
ρ^L	TFP persistence (Low income region)	0.925
$\phi\pi$	Monetary policy inflation coefficient	1.5
ρ_u	Monetary policy shock persistence	0.5

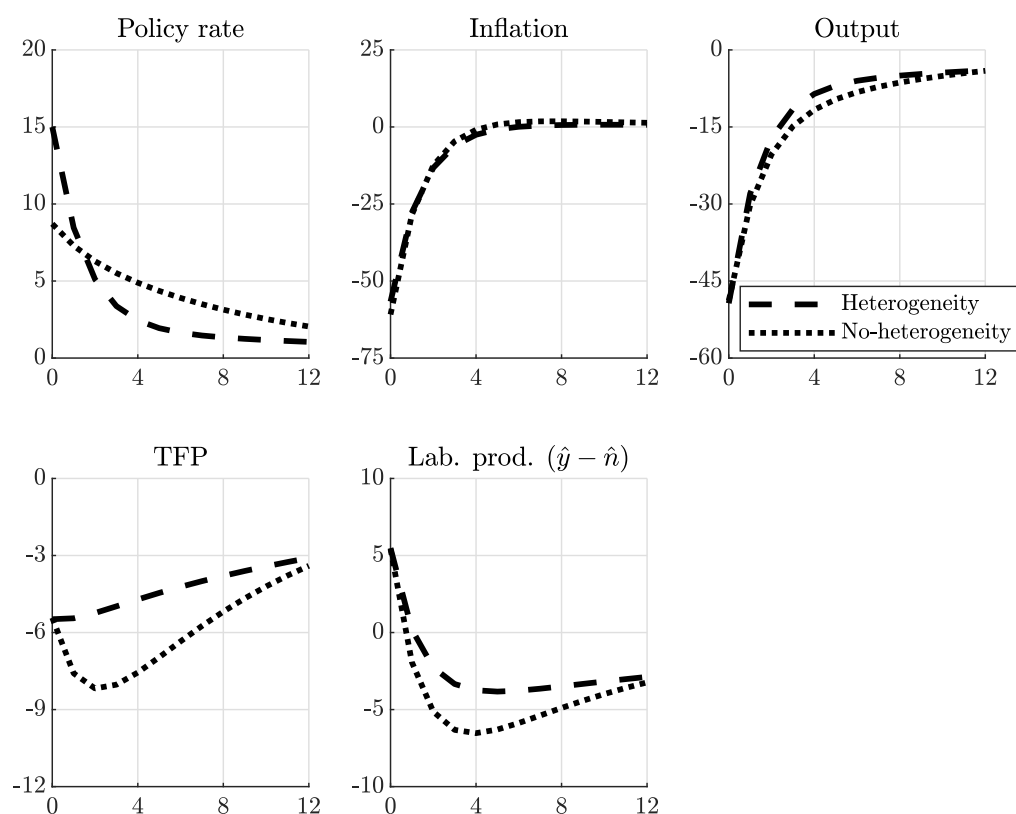
Numerical steady state Given the calibration presented above, the steady state outcomes of the model are presented in Table D.5. Steady state unemployment in the High income region is calibrated to be 5.0%. The consequences of hysteresis is that the Low income region has a higher unemployment rate (at 7.4%) and the income of a High income region is 12.5% higher than a Low income region.

Table D.5: Steady state

Moment	Y^H/Y^L	U^H	U^L
Value	1.125	5.0%	7.4%

Response to a monetary policy shock: Effects of heterogeneity

Figure D.34: Aggregate effects of heterogeneity



Note: 1 percentage point monetary policy tightening. Variables are basis point deviations from steady state. The policy rate and inflation are annualized.

The model also allows us to study the aggregate effects of regional heterogeneity. To this end, we compare in Figure D.34 the previous simulations (dashed lines) with the response to the monetary policy shock when regional heterogeneity is removed (dotted lines). In constructing this counterfactual, we set the steady state unemployment rate in each region to 6.2% and the steady state income level in each region to 0.902, which coincides with the Union-wide steady state values from above. This is achieved by appropriately adjusting ξ and setting the hysteresis parameter to $\rho = 0.877$ for all regions. When heterogeneity across regions is removed, inflation falls a little further on impact and output recovers a little slower. This is because “average ρ ” is lower in the Union with heterogeneity than in the Union with no heterogeneity.

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