Risk Sharing and Monetary Policy Transmission

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Abstract

Using regionally disaggregated data on economic activity, we show that risk sharing plays a key role in shaping the real effects of monetary policy. With weak risk sharing, monetary policy shocks trigger a strong and durable response in output. With strong risk sharing, the response is attenuated, and output reverts to its initial level over the medium term. The attenuating impact of risk sharing via credit and factor markets concentrates over a two-year horizon, whereas fiscal risk sharing operates over longer horizons. Fiscal risk sharing especially benefits poorer regions by shielding them against persistent output contractions after tightening shocks.

Keywords: Monetary Policy; Risk-sharing, Regional Heterogeneity; Local Projections; Quantile Regressions

JEL Classification: C32, E32, E52

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Non-technical summary

This paper provides a novel empirical perspective on the interaction between monetary policy and risk sharing in a currency union. Exploiting regionally disaggregated data for the euro area, we document sizable variation in the overall incidence of risk sharing and in the relative strength of different risk-sharing channels. We then analyse how risk sharing, and its breakdown into fiscal and market-based channels, affect the transmission of monetary policy.

Our estimates point to a key role of risk sharing in shaping the real effects of monetary policy shocks. The regional output response to short-term interest shocks is significantly more pronounced for regions attaining a low degree of risk sharing than for those attaining a high degree. Moreover, regions subject to an elevated degree of risk-sharing are less prone to hysteresis in output.

Fiscal risk sharing proves particularly forceful in shaping the persistence of monetary policy effects in poor regions. With weak fiscal risk sharing, these regions suffer from a durable output contraction in response to a policy-rate hike. By contrast, with strong fiscal risk sharing, poor regions do not only face a weaker output contraction but are also insulated from such hysteresis effects.

Finally, we document some complementarity between private and public risk-sharing channels. Private channels, operating via credit and factor markets, attenuate the real effects of monetary policy over a short time horizon. Public risk sharing, operating via the tax and benefit system, affects monetary policy transmission over longer horizons when private risk-sharing channels loose effectiveness.

These findings provide relevant insights for the debate on the institutional setup of the Economic and Monetary Union (EMU). First, our results suggest that regional heterogeneity in the capacity to absorb shocks via fiscal and market-based channels could interfere with an even transmission of monetary policy. From that point of view, it appears desirable to strive for a harmonised risk-sharing architecture in a currency union. Second, our results point to benefits of risk-sharing in preventing regional hysteresis and related economic divergence. Finally, if risk-sharing tends to dampen the real effect of monetary policy shocks, institutional reform towards more risk sharing could come with the need for a more extensive use of monetary policy to achieve an intended economic stabilisation outcome.

1 Introduction

The literature on optimal currency areas establishes a clear division of labor in the pursuit of macroeconomic stabilization objectives.[1] Monetary policy is to limit fluctuations in average macroeconomic outcomes by adjusting its union-wide stance in response to symmetric shocks. Risk sharing via public and market-based mechanisms instead should limit the dispersion in macroeconomic outcomes across the currency union by facilitating a geographically differentiated adjustment to asymmetric shocks.

An important, but so far under-explored, aspect in implementing this division of labour is that the impact of these macroeconomic stabilization tools may interact. If a given monetary policy stance exerts a uniform impact on different members of a currency union, its role in limiting average economic fluctuations is unaffected by the role of the risk-sharing architecture in limiting their dispersion. But a growing literature has documented that monetary policy transmits unevenly, for instance due to differences in economic structures, initial conditions, and the institutional landscape of the economy.[2] The resultant heterogeneity in the regional incidence of a uniform monetary policy stance may render its overall impact dependent on the risk-sharing architecture of a currency union. For instance, if risk sharing counteracts the stronger contractionary effects of a monetary policy tightening shock in certain regions, also the average contraction may be attenuated compared to a situation without risk sharing.[3]

The current paper provides empirical evidence on the relevance and nature of the interactions between monetary policy and risk sharing. The analysis exploits region-

[1] See, *e.g.*, Mundell (1961); McKinnon (1963); Kenen (1969); and Farhi and Werning (2017).

[2] Monetary-policy relevant aspects of economic structure include, e.g., industry composition and the financing mix of firms (Carlino and Defina, 1998; Peersman and Smets, 2005; Dedola and Lippi, 2005; Hauptmeier et al., 2020). Relevant initial conditions include the state of the business-, credit-, and policy-cycle (Tenreyro and Thwaites, 2016; Jordà et al., 2020; Alpanda et al., 2021; Eichenbaum et al., 2022), as well as the household income and wealth distributions, which may interact with their marginal propensity to consume (Kaplan et al., 2018). Relevant institutional factors include, for instance, the stringency of labor and product market regulations (Christoffel et al., 2009; Blanchard and Gali, 2010; Aghion et al., 2019; Ferrando et al., 2021).

[3] Suppose, *e.g.*, that the initial impact of the shock on disposable income differs and fiscal risk sharing responds by reallocating funds from less to more affected regions. If the net-recipient regions are characterised by a higher marginal propensity to consume than the net-contributing regions, then the overall multiplier on economic activity in the currency union would fall (in absolute terms) as the degree of fiscal risk sharing rises.

ally disaggregated data for the euro area, which exhibit substantial variation in the overall prevalence of risk sharing and in the relative strength of different risk-sharing channels. To estimate the degree and composition of risk sharing, we rely on the well-established framework by Asdrubali et al. (1996). We then feed the resultant coefficients into a standard empirical macro model, estimated via local projections (Jordà, 2005), to assess how risk sharing, and its breakdown into fiscal and market-based channels, shape the transmission of monetary policy. Finally, we apply quantile estimation techniques to explore whether the interaction between monetary policy and risk sharing differs across poorer and richer regions.

On this basis, we derive three main insights. First, risk sharing plays a key role in shaping the real effects of monetary policy shocks. For instance, the regional output contraction after a 100 basis point policy-rate hike is around 1 percentage point shallower for regions attaining the maximum degree of risk sharing in our sample than for those attaining the minimum degree. Moreover, regions subject to a high degree of risk-sharing are less prone to policy-induced hysteresis: while output in regions with minimum risk sharing remains around 1.5 percentage point below its initial level five years after a monetary policy tightening shock, it fully recovers over this period in regions with maximum risk-sharing.

Second, fiscal risk sharing proves particularly forceful in shaping the persistence of monetary policy effects on poor regions. For instance, with weak fiscal risk sharing, these regions suffer from a durable output contraction in response to a policy-rate hike. By contrast, with strong fiscal risk-sharing, poor regions do not only face a weaker output contraction, but are also insulated from such hysteresis effects.

Third, private and public risk-sharing channels act as complements in terms of their respective time profiles. The private channels, operating via credit and factor markets, attenuate the real effects of monetary policy over a horizon of up to two years. Public risk sharing, operating via the fiscal transfer and tax system, instead does not produce significant differences in monetary policy transmission over this period. But it then kicks in over longer horizons, just when the private risk-sharing channels loose effectiveness.

These findings offer relevant insights for the debate on the institutional setup of the Economic and Monetary Union (EMU). First, our results suggest that regional heterogeneity in the capacity to absorb shocks via fiscal and market-based channels could interfere with an even transmission of monetary policy. From that point of view, it appears desirable to strive for a harmonised risk architecture in a currency union. Second, our results point to benefits of risk-sharing in preventing regional hysteresis and related economic divergence. Finally, if risk-sharing tends to dampen the real effect of monetary policy shocks, institutional reform towards more risk-sharing could come with the need for a more extensive use of monetary policy to achieve an intended economic stabilisation outcome.

Related Literature. Our analysis places itself at the intersection of two important strands of the literature. The first relates to the quantification of risk sharing within and across countries (Mace, 1991; Cochrane, 1991; Townsend, 1994). In the seminal approach to this literature, Asdrubali et al. (1996) propose a framework based on decomposing the cross-section of shocks to the GDP of US states to quantify the role of private and public channels risk-sharing channels in smoothing out idiosyncratic shocks. This framework has also been applied to other jurisdictions, including individual euro area countries (e.g. Buettner, 2002; Hepp and von Hagen, 2013 for Germany). While the resultant estimates of intranational risk-sharing lie in a similar range as those for the United States (Burriel et al., 2020), risk-sharing between euro area countries has been found to be significantly lower (Sorensen and Yosha, 1998). At the same time, more recent papers suggest that the amount of international risk sharing in the euro area has risen in the aftermath of the sovereign debt crisis (Cimadomo et al., 2020) and may be further enhanced with deeper fiscal integration (Furceri and Zdzienicka, 2015).

To shed light on how risk sharing interacts with monetary policy transmission, we connect to a second strand of the literature. Starting from the vast body of evidence on the aggregate effects of monetary policy,[4] a growing number of papers have highlighted that the transmission of monetary policy is heterogeneous within a given

[4] See Ramey (2016) for a recent review of that literature.

economy.[5] Applying these insights to the geographical dimension, there is also increasing evidence that the regional incidence of monetary policy differs depending on initial conditions and economic structures (Coeuré, 2018). For instance, Hauptmeier et al. (2020) document that the output response to euro area monetary policy shocks is stronger and more persistent for poorer regions, and similar studies for the US economy point to heterogeneity in transmssion for instance due to differences in local mortgage market conditions (Neville et al., 2012; Fratantoni and Schuh, 2003; Di Maggio et al., 2017; Beraja et al., 2019).

The remainder of the paper proceeds as follows. The next section describes the data and highlights some salient stylized facts regarding economic activity at the regional level. Section 3 presents the estimation of the degree risk-sharing in Euro area countries. The interaction between risk-sharing and monetary policy is analysed in Section 4. Section 5 documents a range of robustness checks and Section 6 concludes.

2 Data and stylized facts

We rely on regionally disaggregated economic data based on Eurostat's Nomenclature of Territorial Units for Statistics (NUTS). The NUTS classification breaks down the EU Member States into four levels. The highest level (NUTS-0) corresponds to the nation state. The lower levels (NUTS-1 to NUTS-3) subdivide national territories into ever more granular units based on population thresholds and existing administrative structures.[6] Our analysis uses NUTS-2 data, which offer the most granular regional breakdown with sufficient variable coverage to estimate the degree of risk-sharing within each country (see Section 3). NUTS-2 regions are defined as hosting between 800,000 and 3,000,000 inhabitants and typically refer to Provinces, Regions and, in some cases, States.[7]

[5] See, *e.g.*, Ampudia Fraile et al. (2018); Coibion et al. (2017); Lenza and Slacalek (2018); Cravino et al. (2020) for papers focusing on the household sector and footnote 2 for papers exploring other relevant dimensions.

[6] See `https://ec.europa.eu/eurostat/web/nuts/principles-and-characteristics`.

[7] For example, German NUTS-2 level regions correspond to the *Regierungsbezirk*. Similarly, in France (until 2015), NUTS-2 level regions correspond to the *Régions*; in Italy to the *Regioni*; in Spain to the *Communidades y ciudades Autónomas*; in the Netherlands to the *Provincies*; and in Austria to

The analysis considers the founding members of the euro area, plus Greece, while excluding Luxembourg and Ireland due to the low number of NUTS-2 regions in the latter two countries.[8] Our final sample thus consists of 155 regions from 10 Euro area countries over the period 2000-2018 at annual frequency.

The main regional variables of interest are gross domestic product (GDP), primary income, disposable income, and consumption. The source of the first three variables is Eurostat's regional statistics database. Regional consumption instead comes from the European Cities and Regions database of Oxford Economics. All variables are deflated to 2015 Euros, using the regional deflator from the European Commission's ARDECO database, and are expressed in per capita terms. We complement this regionally disaggregated information with euro area and country level variables to control for aggregate economic and financial conditions (see Section 4 and Annex C).

The data reveal two key facts that motivate the ensuing analysis. First, there is substantial heterogeneity in economic output across and within euro area countries (Table 1). In 2018, for instance, average per capita GDP at the NUTS-2 level ranged from 14,495€ in Greece to 41,128€ in Austria. Moreover, also within countries, per capita GDP exhibits pronounced dispersion, with some countries displaying a higher coefficient of variation (CV) than the euro area as a whole.

Second, the degree of regional dispersion within countries tends to fall as we turn from GDP to primary income, disposable income, and ultimately consumption. On average, the regional CV for GDP is 2.5 times as high as that for consumption. By contrast, the corresponding figure amounts to only 1.3 for the euro area. This pattern points to powerful forces that weaken the link between regional output and consumption spending within countries. In the following section, we shed light on these forces through the lens of their risk-sharing properties, before exploring their interaction with monetary policy transmission in Section 4.

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the *Bundesländer*.

[8]Luxembourg and Ireland contain only 1 and 3 NUTS-2 regions, respectively. The sample for Greece starts in 2001 because this is when the country introduced the euro. In 2014, the French parliament passed a law reducing the number of metropolitan regions from 22 to 13. We conserve the former territorial division of 22 NUTS-2 regions. We also follow the literature in excluding NUTS-2 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).

Table 1: Descriptive statistics

	GDP		Prim. income		Disp. income		Consumption	
	Mean	*CV*	*Mean*	*CV*	*Mean*	*CV*	*Mean*	*CV*
Austria	41,128	17	26,628	5	22,545	3	20,860	2
Belgium	36,521	35	24,708	14	19,918	9	19,255	6
Finland	40,929	20	24,237	16	21,205	9	21,590	11
France	29,859	23	21,145	12	19,047	6	17,568	4
Germany	37,345	23	26,398	16	21,174	9	19,508	7
Greece	14,494	22	9,988	17	9,774	12	11,127	8
Italy	28,190	29	18,612	26	17,088	20	17,272	20
Netherlands	39,254	23	25,209	10	18,866	5	18,324	4
Portugal	18,507	19	11,339	16	11,619	13	11,824	9
Spain	23,982	21	15,273	20	14,245	17	13,952	18
Euro area	29,664	27	20,888	30	17,727	24	17,225	20

Note: Figures refer to real per capita GDP, primary income, disposable income and consumption in 2018 at the NUTS-2 level, except for the euro area row, which is based on NUTS-0 (country-level) data. The coefficient of variation (CV) is computed as the ratio of the standard deviation to the mean of all NUTS-2 (NUTS-0) units within each country (the euro area) in 2018.

3 Risk-sharing in euro area countries

Following Asdrubali et al. (1996), we identify three channels of risk-sharing: the factor market channel, the fiscal channel, and the credit market channel.[9] To this end, we estimate the following panel regressions for each euro area country at the NUTS-2 level, using annual data from 2000 to 2018:

$$\Delta^j gdp_t^k - \Delta^j pi_t^k = \beta_{K,j} \times \Delta^j gdp_t^k + \alpha_{K,t} + \varepsilon_{K,t}^k \quad \textit{(Factor market channel)} \tag{1}$$

$$\Delta^j pi_t^k - \Delta^j di_t^k = \beta_{F,j} \times \Delta^j gdp_t^k + \alpha_{F,t} + \varepsilon_{F,t}^k \quad \textit{(Fiscal channel)} \tag{2}$$

$$\Delta^j di_t^k - \Delta^j c_t^k = \beta_{C,j}^c \times \Delta^j gdp_t^k + \alpha_{C,t} + \varepsilon_{C,t}^k \quad \textit{(Credit market channel)} \tag{3}$$

$$\Delta^j c_t^k = \beta_{U,j} \times \Delta^j gdp_t^k + \alpha_{U,t} + \varepsilon_{U,t}^k \quad \textit{(Unsmoothed)} \tag{4}$$

[9] See Annex A for additional detail on the methodology. We use the terms fiscal and public risk-sharing as synonyms and occasionally refer to the factor and credit market channels as private risk-sharing.

where gdp_t^k denotes the log of real per capita output in region k in period t, and pi_t^k, di_t^k and c_t^k are the corresponding variables for primary income, disposable income, and consumption of private households, respectively. $\alpha_{x,t}$ are time fixed-effects and $\varepsilon_{x,t}^k$ are the error terms. As the degree of risk-sharing through either of the channels may vary with the persistence of the underlying GDP fluctuations, we estimate these equations over alternative time-differencing intervals, such that $\Delta^j x_t^k = x_t^k - x_{t-1-j}^k$, with $j = 0, ..., 5$ (Asdrubali et al., 1996; Athanasoulis and van Wincoop, 2001). The equations are estimated via OLS, with standard errors clustered at the regional level and boostrapped using 1000 iterations.

The wedge between output and primary income corresponds to the net income streams receivable from and payable to other regions and countries; these may originate either from capital, due to internationally or inter-regionally diversified private investment portfolios (*e.g.* yielding dividend payments); or from labour, due to the compensation accruing to commuters who work in a region or country other than their place of residence.[10] Accordingly, $\beta_{K,j}$ measures the amount of risk-sharing achieved through the factor market channel.

The wedge between primary and disposable income stems from the difference between tax payments to and transfer payments from the government; $\beta_{F,j}$ thus measures the fraction of shocks smoothed by the fiscal channel.

Finally, the wedge between disposable income and consumption reflects economic agents' debt accumulation minus savings in each period, so that $\beta_{C,j}$ measures the inter-temporal risk-sharing via credit markets. The amount of unshared risk in the economy, in turn, is given by $\beta_{U,j}$, so that the β-coefficients jointly satisfy:

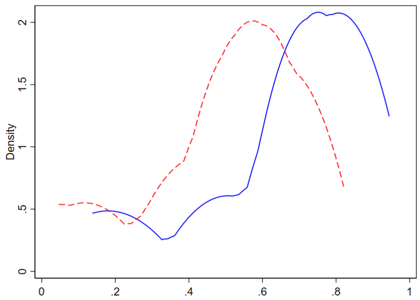
$$\beta_{K,j} + \beta_{F,j} + \beta_{C,j} = \beta_{S,j} = 1 - \beta_{U,j} \tag{5}$$

[10] The wedge between output and primary income is analogous to the difference between GDP and GNP in national accounting. Given the geographical proximity between regions, labour income is likely to account for a larger share of this wedge at the sub-national than at the cross-country level. But also within countries, the size and composition of this wedge may differ noticeably between regions, also due to diverse commuting patterns. In 2018, for instance, the ratio of primary income to GDP was particularly low in capital regions, such as Région de Bruxelles-Capitale/Brussels Hoofdstedelijk Gewest (34.1%) in Belgium. By contrast, this ratio was significantly larger in other regions, such as Schleswig-Holstein in Germany (83.0%) or Burgenland (87.6%) in Austria.

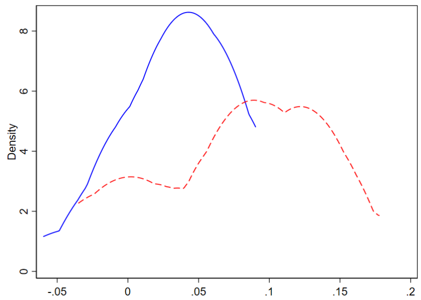
where $\beta_{S,j}$ is the total amount of risk-sharing in a given region.

The estimates of equations (1)-(4) document substantial cross-country heterogeneity in risk-sharing patterns (Figure 1 and Annex Tables B1-B2). Factor markets generally emerge as the dominant channel in smoothing out contemporaneous fluctuations in GDP (*i.e.* setting $j = 0$): the respective point estimates, $\hat{\beta}_{K,0}$ are highly significant and exceed $\hat{\beta}_{F,0}$ and $\hat{\beta}_{C,0}$ by a substantial margin in most countries.[11] At the same time, the cross-country distribution of $\hat{\beta}_{K,0}$ covers a broad range, which we exploit in Section 4 to study the interaction between varying risk-sharing intensities and monetary policy transmission. Likewise, the estimates for the other two channels exhibit pronounced cross-country heterogeneity, while presenting a more uneven picture in terms of relative strength and precision.

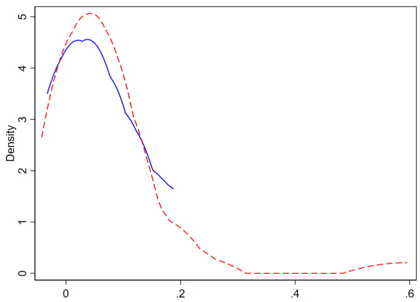
Figure 1: Cross-country densities of estimated risk-sharing coefficients



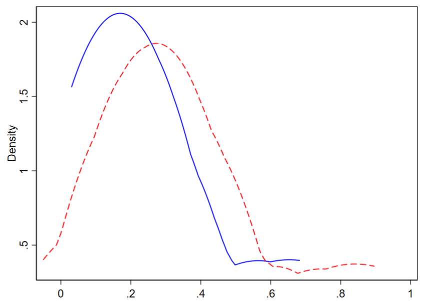
(a) Factor market channel



(b) Fiscal channel



(c) Credit market channel



(d) Unsmoothed

Note: Blue solid (red dashed) lines show the cross-country density functions of the β-coefficients estimated in equations (1)-(4) for $j = 0$ ($j = 5$). The former (latter) correspond to the amount of risk-sharing achieved contemporaneously (over a five-year horizon). The full set of estimates is reported in Tables B1-B2.

[11] The prominent risk-sharing contribution of factor markets is consistent with previous studies. For instance, Hepp and von Hagen (2013) document that factor markets accounted for more than 50% of risk-sharing across German States in the post-reunification period.

The persistence of the regional GDP fluctuations proves particularly relevant for the factor market and fiscal channels. As the time-differencing interval expands, $\beta_{K,j}$ declines and, after five years, almost the entire cross-country distribution lies to the left of that for the contemporaneous coefficients (see red dashed and blue solid lines in Figure 1a). By contrast, fiscal risk-sharing tends to intensify for higher j's, albeit insufficiently so to prevent a rise in the unsmoothed portion (Figures 1b and 1d); and, in most countries, $\beta_{F,j}$ turns significant over longer horizons. Consistent with the findings in Asdrubali et al. (1996), this pattern points to a sluggish response of the smoothing mechanisms working via the government sector.

4 Risk-sharing and monetary policy transmission

In this section, we study how the degree of risk-sharing in an economy affects the transmission of monetary policy to the real economy. To this end, we first set up a benchmark local linear projections model to estimate the dynamic effect of monetary policy on regional output. We then augment this model with the estimated parameters reported in Section 3 to assess whether risk-sharing reinforces or dampens the real effects of monetary policy. Third, we break down the impact of risk-sharing into its private and public channels. Finally, we study whether the impact of specific risk-sharing channels differs across poorer and richer regions.

4.1 Baseline model

As a starting point, we follow Hauptmeier et al. (2020) in estimating 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:

$$y_{k,t+h} = \alpha_{k,h} + \kappa_h i_t + \gamma_h \mathbf{X}_{k,t} + \delta_h \mathbf{X}_{c,t} + \theta_h \mathbf{X}_t + \varepsilon_{k,t+h} \tag{6}$$

where the dependent variable $y_{k,t+h}$ denotes real GDP (in log) in jurisdiction k and year $t+h$; α_k is a set of region-fixed effects; i_t is the monetary policy-controlled short-term interest rate in year t and enters equation 6 as a percentage per annum;[12] $\mathbf{X_{k,t}}$, $\mathbf{X_{c,t}}$, and $\mathbf{X_t}$ are vectors of time-variant control variables at the region-, country- and euro area-level, respectively (see Table C1); and $\varepsilon_{k,t+h}$ is an error term.

At the regional level, we control for the population (in log) of each NUTS-2 unit. At the country level, we control for: *(i)* financial conditions, captured by stock market indices (in log) and the 10-year government bond yield; and *(ii)* fiscal positions, captured by the government debt ratio and the change in the structural primary balance. The euro area controls include GDP and HICP (both in log) as standard elements of the monetary policy reaction function.

Accordingly, our main identification assumption is that, controlling for aggregate macroeconomic conditions, monetary policy is exogenous to variation in regional GDP. As discussed in Hauptmeier et al. (2020), this assumption is supported by: *(i)* the ECB's mandate pertaining to the euro area level (ECB, 1998); *(ii)* the small size of individual regions relative to the euro area economy as a whole;[13] and *(iii)* the long publication lags for regional data, which become available only around two to three years after the period they refer to and, as such, are not part of the policy-relevant information set in real-time. In Section 5 we also subject this identification assumption to a range of robustness checks.

We estimate equation 6 by ordinary least squares (OLS) for $h = 0, \ldots, 5$. Inference is based on Driscoll and Kraay (1998) standard errors that account for cross-sectional and temporal dependencies in the data. The results are presented as the response of regional output to a 100 basis point hike in the short-term interest rate in year t, captured by the coefficient κ_h for each horizon h.

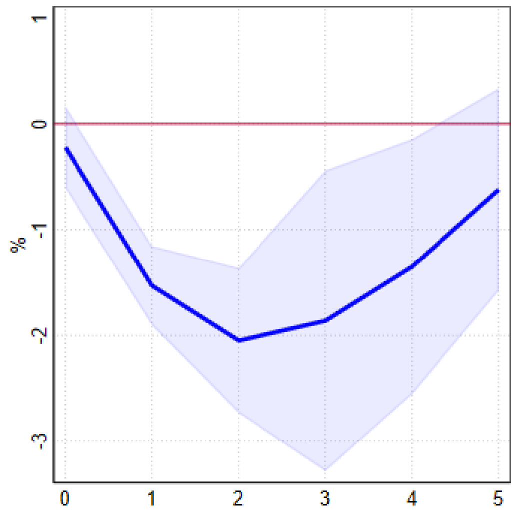
As depicted by the impulse response function (IRF) in Figure 1, regional output drops in the first year by around 0.2%. The downward impact on economic activity further amplifies and reaches its peak at the two-year horizon, with the impulse re-

[12] Following the literature, we use the 3-month Euribor as our measure of the policy-controlled short-term interest rate (e.g. Smets and Wouters, 2003). The data for this variable come from the Area Wide Model (AWM) database. To better capture the impact of non-standard measures affecting the policy stance in the final years of the estimation period, we extend the short-term interest rate from 2014 by adding the cumulative changes of the shadow interest rate developed by Lemke and Vladu (2017).

[13] For example, while Lombardia's GDP accounted for 22% of Italian GDP in 2018, it represented only 3.5% of Euro area GDP.

sponse function implying a contraction of around 2% in regional GDP in year $t+2$. Output then gradually recovers over the remainder of the IRF horizon. These estimates are consistent with the related literature on the real effects of monetary policy (e.g. Smets and Wouters, 2003) and with previous studies using regional data for the euro area (e.g. Hauptmeier et al., 2020).

Figure 1: Impact of monetary policy on regional aggregates



Note: 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.

4.2 Risk-sharing and monetary policy transmission

We now turn to the question on how risk-sharing interacts with monetary policy transmission. To this end, we augment the baseline model (equation 6) with the risk-sharing estimates from Section 3, which yields:

$$y_{k,t+h} = \alpha_{k,h} + \left(\kappa_{0,h} + \kappa_{S,h} \times \hat{\beta}^{c}_{S,h}\right) i_t + \gamma_h \mathbf{X}_{k,t} + \delta_h \mathbf{X}_{c,t} + \theta_h \mathbf{X}_t + \varepsilon_{k,t+h} \tag{7}$$

where $\hat{\beta}^{c}_{S,h}$ is the amount of risk-sharing achieved in country c after h periods. Because they are estimated from the data, the $\hat{\beta}^{c}_{S,h}$ coefficients are generated regressors and have their own sampling variance (Pagan, 1984; Murphy and Topel, 1985). We thus follow the literature and bootstrap both stages of the analysis (*i.e.* the estimation of risk-sharing in section 3 and the estimation of the local projections in section 4.2) to adjust the standard errors. The Driscoll and Kraay (1998) standard errors are therefore bootstraped using 1000 iterations. We use the same control variables as in the baseline model.

To facilitate the interpretation, we standardize the $\hat{\beta}_{S,h}$ coefficients for each country c by demeaning and dividing it by its cross-country standard deviation. Thus, the coefficient $\kappa_{0,h}$ captures the response of regional output in period $t+h$ to a change in the short-term interest rate in period t when risk-sharing is at the average of the sample. The coefficient $\kappa_{S,h}$ shows whether and how this impact varies with the degree of risk sharing. Taken together, these coefficients summarize the impact of a 100 basis point monetary policy rate hike at each horizon h conditional on the degree of risk-sharing in country c as:

$$\frac{\partial y_{k,t+h}}{\partial i_t} = \kappa_{0,h} + \kappa_{S,h} \times \hat{\beta}^c_{S,h} \tag{8}$$

We report the estimation of equation 7 in Table 2. The interaction term between the monetary policy interest rate and risk-sharing is positive and statistically significant for the entire horizon. Hence, our results indicate that risk sharing dampens the real effects of monetary policy. To highlight the economic relevance of this dampening effect, Figure 2 presents the response of regional output to a 100 basis point interest rate hike for different percentiles of the risk-sharing distribution. At the upper quartile of the distribution, regional output decreases by 1.9 % after 2 years. This is 0.4 percentage point higher than when risk-sharing is at the lower quartile of the distribution (Figure 2a).

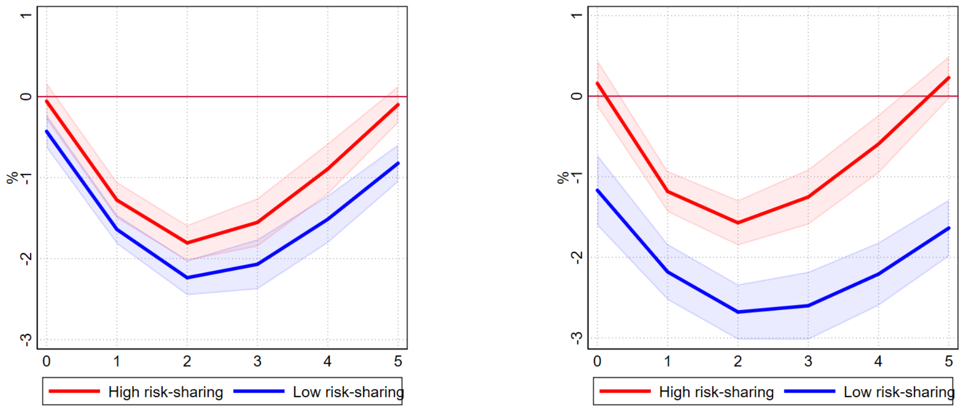
Further, these differences in the output response intensify when considering the outer parts of the risk-sharing distribution. For instance, the GDP trough for the lower decile (again materialising in year $t+2$) is 1.1 percentage points lower than the corresponding trough for the upper decile (Figure 2b). Moreover, the persistence of the GDP contraction differs markedly across the distribution. At high degrees of risk-sharing, output recovers to its pre-shock levels by the end of the horizon. At low degrees of risk-sharing, the contractionary effect of monetary policy instead proves

Table 2: Baseline estimates for coefficients on the short-term interest rate and the interaction with risk-sharing.

	$h=0$	$h=1$	$h=2$	$h=3$	$h=4$	$h=5$
i_t	-0.331***	-1.593***	-2.092***	-1.853***	-1.279***	-0.556***
	(0.121)	(0.111)	(0.124)	(0.181)	(0.183)	(0.134)
$i_t \times \hat{\beta}^c_{S,h}$	0.486***	0.364***	0.387***	0.477***	0.582***	0.681***
	(0.136)	(0.115)	(0.108)	(0.109)	(0.108)	(0.101)
Observations	2945	2790	2635	2480	2325	2170
Within R2	0.705	0.698	0.663	0.595	0.529	0.514
Number of regions	155	155	155	155	155	155

Note: This table reports the estimation of equation 7. We do not report the estimations of the control variables. $\hat{\beta}_S$ are standardized. The Driscoll and Kraay (1998) standard errors are given in parenthesis. Standard errors are bootstrapped using 1000 interactions. * / ** / *** indicate 1% / 5% / 10% significance level.

Figure 2: Impact of monetary policy on regional output when risk is shared



(a) Upper versus lower quartiles

(b) Upper versus lower deciles

Note: 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. Red (blue) lines depict the estimates for the upper (lower) quartiles or deciles of $\hat{\beta}^c_{S,h}$. The Driscoll and Kraay (1998) standard errors are boostrapped using 1000 iterations.

persistent and output remains depressed even five years after the shock. In summary, by smoothing out the output losses in individual regions, risk sharing also weakens the average effects of monetary policy on the economy and renders GDP less susceptible to hysteresis effects (Blanchard, 2018).

4.3 Interrelation of private and public risk-sharing

We next examine the contribution of individual risk-sharing channels in shaping monetary policy transmission. We hence replace $\hat{\beta}^c_{S,h}$ by its components, as listed in equation 5, to estimate the following model:

$$
\begin{aligned}
y_{k,t+h} = &\alpha_{k,h} + \left(\kappa_{0,h} + \kappa_{K,h} \times \hat{\beta}^c_{K,h} + \kappa_{F,h} \times \hat{\beta}^c_{F,h} + \kappa_{C,h} \times \hat{\beta}^c_{C,h}\right) i_t \\
&+ \gamma_h \mathbf{X}_{k,t} + \delta_h \mathbf{X}_{c,t} + \theta_h \mathbf{X}_t + \varepsilon_{k,t+h}
\end{aligned} \tag{9}
$$

where $\hat{\beta}^c_{K,h}$, $\hat{\beta}^c_{F,h}$, $\hat{\beta}^c_{C,h}$ are the amount of risk-sharing achieved in country c after h periods through the factor market channel, the fiscal channel, and the credit market channel, respectively. As in equation 7, the β-coefficients are standardized and the results refer to the impact of a 100 basis point rate hike. The control variables are the same as in regressions 6 and 7 and the Driscoll and Kraay (1998) standard errors are again bootstrapped using 1000 iterations.

Table 3: Baseline estimates for coefficients on the short-term interest rate and the interaction with the fraction of shared risk.

	$h=0$	$h=1$	$h=2$	$h=3$	$h=4$	$h=5$
i_t	-0.106	-1.290***	-2.051***	-1.842***	-1.272***	-0.562***
	(0.143)	(0.145)	(0.156)	(0.202)	(0.206)	(0.159)
$i_t \times \hat{\beta}^h_{K,h}$	1.177***	0.619***	0.348***	0.337***	0.297**	0.269**
	(0.192)	(0.164)	(0.131)	(0.124)	(0.132)	(0.124)
$i_t \times \hat{\beta}^h_{F,h}$	-0.0335	0.114	0.536***	0.648***	0.735***	0.759***
	(0.169)	(0.181)	(0.117)	(0.121)	(0.110)	(0.107)
$i_t \times \hat{\beta}^h_{C,h}$	1.293***	0.813***	0.396**	0.491***	0.659***	0.716***
	(0.216)	(0.234)	(0.177)	(0.137)	(0.135)	(0.135)
Observations	2945	2790	2635	2480	2325	2170
Within R2	0.735	0.716	0.680	0.613	0.550	0.533
Number of regions	155	155	155	155	155	155

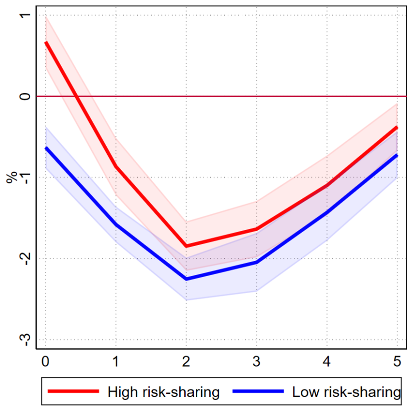
Note: This table reports the estimation of equation 9. We do not report the estimations of the control variables. $\hat{\beta}^c_{K,h}$, $\hat{\beta}^c_{F,h}$, $\hat{\beta}^c_{C,h}$ are standardized. The Driscoll and Kraay (1998) standard errors are given in parenthesis. Standard errors are bootstrapped using 1000 interactions. * / ** / *** indicate 1% / 5% / 10% significance level.

The significant and positive interaction terms imply that both, private risk-sharing via factor and credit markets, as well as public risk-sharing via fiscal arrangements, cushion the impact of a monetary policy tightening (Table 3). However, the channels differ in their time profiles. Private risk-sharing channels tend to dampen the monetary policy shock contemporaneously and up to one year after the shock. For example, Figure 3a (Figure 3c) shows that regional output drops by around 0.9% (0.8%) after one year when the degree of risk-sharing through factor (credit) markets is in the upper quartile of the distribution. Instead, the contraction is significantly more pronounced, reaching close to 1.6% (2.1%), when risk-sharing through these channels is in the lower quartile. After two years, however, the dampening impact of risk-sharing through factor and credit markets fades away.

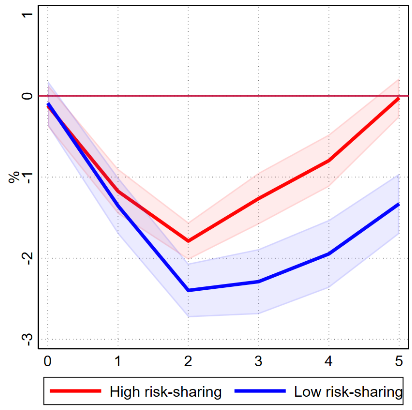
Fiscal risk-sharing instead tends to dampen the economic consequences of a rate hike over longer horizons (Figure 3b). While differences in the point estimates are not statistically significant up to $t+1$, the IRFs as well as the confidence bands diverge as of $t+2$. When fiscal risk-sharing is in the upper quartile, regional output drops by slightly less than 1.8% in year $t+2$ and recovers to its initial level after five years. By contrast, the estimates point to a sharper contraction when fiscal risk-sharing is in the lower quartile, reaching 2.4% after two years. Moreover, at the end of the horizon, output remains depressed, still standing 1.3% below its level prior to the shock in economies with a low amount of fiscal risk-sharing.

The diverse time profiles across private and public channels appear consistent with the following interpretation, which echoes similar considerations by Asdrubali et al. (1996). Credit markets offer a well-established and timely avenue for private agents to smooth out fluctuations in income over the economic cycle, *e.g.* by borrowing from banks. However, as downturns become more persistent, banks will gradually pare back their lending activity due to declining creditworthiness, which in turn lowers the respective risk-sharing contribution over time. Similarly, cross-regional income flows would initially be unaffected by a downturn in a specific economy, thus supporting the factor market channel in the initial years after the shock. But as the downturn drags on, households may be forced to divest their international asset holdings, for instance, thus foregoing the respective future income streams. As a consequence, also the the

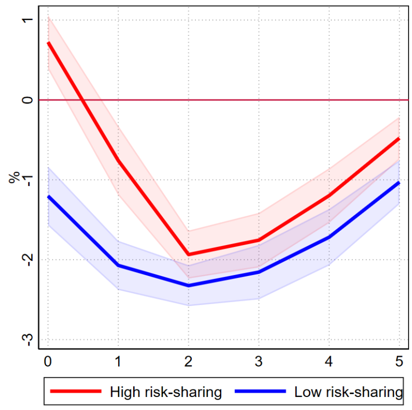
Figure 3: Impact of monetary policy on regional output when risk is shared through the different channels



(a) Factor market channel



(b) Fiscal channel



(c) Credit market channel

Note: 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. Red (blue) lines depict the estimates for the upper (lower) quartiles of $\beta^{c}_{K,h}$, $\beta^{c}_{F,h}$ or $\beta^{c}_{C,h}$. The Driscoll and Kraay (1998) standard errors are boostrapped using 1000 iterations.

factor market channel loses potency as the downturn persists.

The gradual phasing-in of public risk-sharing reflects the following factors. In line with the findings of Asdrubali et al. (1996) and Buettner (2002), our empirical estimates suggest lagged effects of fiscal transfers (see Figure 1b and Table B1), which are consistent with the usual lags in the fiscal response to changing economic circumstances. These lags tend to be particularly pronounced for discretionary fiscal policies, which need to first garner political approval and, even thereafter, often require significant lead time until implementation. But also automatic stabilizers may operate in a sluggish manner, related for example to specificities of the tax code implying a lagged collection of revenues (e.g. Bouabdallah et al., 2020); or to labour market institutions shaping the transmission of economic shocks to labour markets (e.g. Abbritti and Weber, 2018) and therefore the timing of unemployment-related benefit payments (Buettner, 2002). As a result, the fiscal risk-sharing channel shows up significant only from year $t+2$ onward. But it then leads to a materially faster recovery and prevents the persistent contraction in output arising in regions with a low level of fiscal risk-sharing.

Overall, our estimates thus point to complementarities between private and public risk-sharing channels over time. While private risk-sharing tends to dampen the effects of a monetary policy shock in the short-run, public risk-sharing tends to limit adverse regional output effects over longer horizons and to prevent long-lasting hysteresis effects. In the next section, we ask whether risk-sharing mechanisms operate with the same intensity between the poorer and richer regions.

4.4 Heterogeneity across regions

We next explore whether the interaction of risk-sharing with the transmission of monetary policy varies between poor and rich regions. Given its explicit redistributive character, we mainly focus on the fiscal risk-sharing channel. Even without inequality across economic agents, fiscal instruments may attenuate disposable income fluctuations and thereby stabilize consumption and output (Brown, 1955). But the stabilization role of fiscal policy may be reinforced in the presence of heterogeneity, if net transfers are targeted towards agents with a larger marginal propensity to spend (Blinder, 1975), which tend to concentrate in the lower parts of the income distribution (Parker et al., 2011). As documented by McKay and Reis (2016) for the US economy, the stabilizing properties of fiscal policy indeed mainly derive from its effects in mitigating household inequality and its social insurance function. As poorer geographical units tend to host a larger share of vulnerable households and firms (Hauptmeier et al., 2020), these mechanisms may also operate at the regional level.

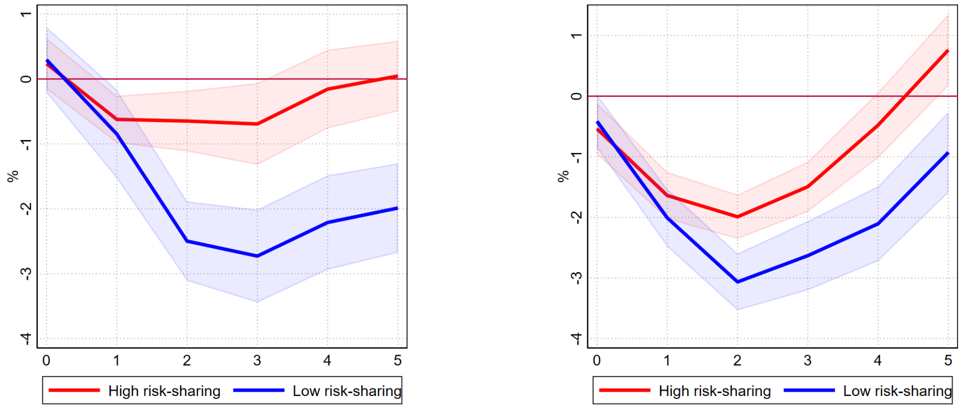
To quantify the dynamic impact of exogenous changes in monetary policy across the regional GDP distribution for different levels of risk-sharing, we combine Jordà (2005)'s local projections method with quantile estimation techniques. First introduced by Koenker and Bassett (1978), quantile regression models characterize the entire conditional distribution of a dependent variable conditional on a set of regressors. These models therefore also provide a flexible way to explore heterogeneity in the response to monetary policy and its interaction with risk sharing. In the presence of fixed effects, quantile estimation suffers from incidental parameter problems (Lancaster, 2000). To address this issue, we employ the quantiles-via-moments estimator proposed by Machado and Santos Silva (2019) to estimate panel data models with individual fixed effects. This approach enables us to control for unobserved heterogeneity while estimating quantile-specific coefficients of the covariates in our model via location- and scale-functions.[14]

The quantile estimates reveal pronounced differences in the degree to which fiscal risk sharing shapes the transmission of monetary policy to rich versus poor regions (defined here as the upper and lower decile of the conditional GDP distribution). With weak fiscal risk-sharing, GDP in poor regions does not only exhibit a strong contraction, but the impact proves highly persistent (Figure 4a): five years after the shock, GDP still remains around 2% below its initial level.[15] By contrast, with strong fiscal risk-sharing, the GDP contraction in poor regions is markedly shallower and turns insignificant at horizon $t+4$. For rich regions, stronger fiscal risk-sharing also dampens the real effects of monetary policy (Figure 4b). But the difference in the strength and, especially, the persistence of these effects across risk-sharing levels is much less accentuated than for poor regions. As such, fiscal risk-sharing emerges as particularly instrumental in preempting long-lived 'hysteresis' effects of monetary policy in

[14] In the quantile regressions, we again account for potential error correlation across space and time. To this end, we resort on two-way clustering, given the Discoll-Kraay correction used for the mean regressions is not available for the quantile estimator. As highlighted by Machado and Santos Silva (2019), the estimator might be biased when the number of cross-sectional units is large relative to the the time periods. As a robustness check, we thus use the split-sample jackknife bias correction of Dhaene and Jochmans (2015) to address this potential issue and confirm that our results remain intact.

[15] This evidence on long-lived real effects of monetary policy in poor regions echoes similar findings for an even more granular geographical breakdown in Hauptmeier et al. (2020).

Figure 4: Impact of monetary policy on regional output of poor and rich regions with high or low levels of fiscal risk-sharing



(a) Effect on poor regions (b) Effect on rich regions

Note: 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. Red (blue) lines depict the estimates for the upper (lower) deciles of $\hat{\beta}^{c}_{F,h}$ in poor (10th percentile) and rich (90th percentile) regions. Standard errors are clustered at the time and regional-level and are boostrapped using 1000 iterations.

regions with weak economic performance already prior to the shock.[16]

[16] As an interesting aside, the factor- and credit-market channels, if anything, operate more forcefully in the upper part of the GDP distribution (Figure 5). Although the patterns are not as clear-cut as for fiscal risk-sharing, this result appears consistent with the non-redistributive nature of these two channels and with the notion that the ability to build diversified asset portfolios and to maintain easy access to credit markets concentrates in the higher quantiles of the income distribution.

5 Robustness

As a first robustness check, we rerun our baseline estimations without the 20 largest regions for each year in our sample (Figure 6 in Annex E.1). This allows us to account for the possibility that monetary policy decisions are not entirely independent of economic conditions in regions that carry a particularly large weight in aggregate GDP. The estimated coefficients remain very close to our baseline results, thus supporting our initial assumption that regional conditions, even for the largest regions, do not enter the central bank objective function.

In a second robustness check, we further systematize this evidence by explicitly controlling for regional heterogeneity in the local projections model. Recent evidence for the United States has highlighted that the Federal Reserve indeed reacts to regional disparities in its interest-rate setting (Coibion and Goldstein, 2012). If confirmed also for the euro area, these regularities would cast doubt on our main identifying assumption that, controlling for aggregate economic conditions, monetary policy is exogenous to regional GDP. We hence calculate a set of regional economic dispersion measures and test whether they exhibit significant explanatory power in a standard monetary policy reaction function. Following Coibion and Goldstein (2012), the dispersion measures comprise: *(i)* the population-weighted variance of per-capita regional GDP, *(ii)* the difference between GDP of the 90th and 10th percentiles of the cross-sectional distribution of regional output at time t and *(iii)* a similar measure but for the 80th and 20th percentiles. None of these dispersion measures carries a significant coefficient in the estimated policy reaction function.[17] Since the absence of significan coefficients may also reflect the low number of observations in our time sample (amounting to just 18 data points), we also conduct additional exercises including each of the dispersion measures as an additional covariate in the local projections model. Again, this modification to our baseline model leaves our results intact.[18]

Third, we check the robustness of our results to the use of different shadow rate

[17] As in Coibion and Goldstein (2012), the reaction function defines the short-term interest rate as the dependent variable and, as explanatory variables, includes the first lag of the dependent variable, along with euro area GDP growth and inflation, as well as the first lag of the respective dispersion measures.

[18] Impulse response functions are available upon request.

measures. The pronounced model uncertainty in deriving shadow rates has given rise to a large dispersion of estimates available in the literature (Figure 7 in Annex E.2). We hence replace our baseline measure, as described in footnote 12, with two alternatives, namely the ones developed by Krippner (2015) and by Wu and Xia (2017). Again, our results are robust to these alternative options (Figure 8 in Annex E.2).

To guard against omitted variable bias, we extend our baseline model with a set of factors that may be correlated with both regional activity and policy rates. Specifically, we include global oil prices and the real effective exchange rate against major trading partners as they: *(i)* often feature in ECB policy communication; and *(ii)* are likely to have different effects on each region's economic activity depending on their respective economic structures (e.g. House et al., 2020). The inclusion of these two variables yield impulse response functions close to our baseline estimations (Figure 9, see Annex E.3). Moreover, we run further checks with a modified set of control variables. In particular, we replace the long-term government bond yield by its spread vis-à-vis Germany to capture country-specific risk premia, which played a prominent role during the euro area sovereign debt crisis. We also use an alternative measure for the fiscal stance by replacing the change in the structural primary balance with the change in the overall government balance. As reported in Figure 10 (Annex E.4), the resultant estimates confirm our baseline estimates.

Finally, monetary policy decisions may not only react to current macroeconomic conditions but also to the expected realizations of economic activity and prices. To this end, we rely on the quarterly ECB staff macroeconomic projections to construct forward-looking versions of these variables. To this end, we compute the average of the two-year projections across the four projection vintages of each year (see Appendix E.5 for further detail). Since medium-term macroeconomic projections are highly correlated with current economic developments, we replace the contemporaneous variables for euro area GDP and prices by its medium-term projections to prevent from multicollinearity issues. Again, the estimates remain largely unchanged compared with our benchmark results (Figure 11).

6 Conclusion

This paper explores the interaction between risk sharing and monetary policy exploiting granular data on economic activity in euro area regions. Our estimates show that risk sharing varies across euro area countries, channels, and with the persistence of shocks. This variation has implications for the transmission of monetary policy. First, we find that risk sharing shapes the real effects of monetary policy shocks. Regions attaining a high degree of risk sharing experience a significantly shallower contraction in output after a policy-rate hike and are less prone to policy-induced hysteresis. Second, public risk sharing especially benefits poor regions by shielding then against hysteresis effects. Third, fiscal and market-based risk-sharing channels emerge as complements in that they operate at different time horizons.

These findings speak to an active policy debate on the merits of deeper fiscal and financial integration in the euro area (Bénassy-Quéré et al., 2018). In this context, it has been argued that increased risk-sharing would strengthen the euro area's resilience to economic shocks (Draghi, 2018; Lane, 2021). Our empirical analysis provides support to this notion in that it highlights the benefit of risk sharing in mitigating adverse hysteresis effects in policy-tightening cycles. At the same time, it also indicates that the degree and composition of risk sharing has a bearing on how strongly monetary policy needs to be adjusted in order to achieve a given effect on economic activity.

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A Methodology to estimate the degree of risk-sharing

We follow Asdrubali et al. (1996) and decompose the cross-sectional variance of shocks to GDP by first considering the following identity, holding for any period t:

$$GDP^k = \frac{GDP^k}{PI^k}\frac{PI^k}{DI^k}\frac{DI^k}{C^k}C^k \tag{10}$$

where all the magnitudes are in per capita terms and k is an index of regions. To emphasize the cross-sectional nature of our derivation, we abstract from the time index. GDP denotes regional gross domestic product. PI, DI and C denote, for private households, primary regional income, disposable regional income and regional consumption respectively.

Taking logs and differences of the identity 10, multiply both sides by $\Delta \log GDP$, and take the cross-sectional average to obtain the variance decomposition:

$$\begin{aligned}\text{var}\{\Delta \log GDP\} =& \text{cov}\{\Delta \log GDP, \Delta \log GDP - \Delta \log PI\} \\ &+ \text{cov}\{\Delta \log GDP, \Delta \log PI - \Delta \log DI\} \\ &+ \text{cov}\{\Delta \log GDP, \Delta \log DI - \Delta \log C\} \\ &+ \text{cov}\{\Delta \log GDP, \Delta \log C\} \end{aligned} \tag{11}$$

Dividing by $\text{var}\{\Delta \log GDP\}$ yields the following equation:

$$\beta_K + \beta_F + \beta_C = 1 - \beta_U \tag{12}$$

which corresponds to equation 5 for $j = 0$, where, for example:

$$\beta_K = \text{cov}\{\Delta \log GDP, \Delta \log GDP - \Delta \log PI\} / \text{var}\{\Delta \log GDP\}$$

β_K is the ordinary least squares estimate of the slope in the cross-sectional regression of $\Delta \log GDP$ on $\Delta \log GDP - \Delta \log PI$ (equation 1 for $j = 0$) , and similarly for β_F (equation 2), β_C (equation 3) , and β_U (equation 4).

B Risk-sharing estimation results

Table B1: Estimation of the $\hat{\beta}_K$ and $\hat{\beta}_F$-coefficients

	$\hat{\beta}_{K,j}$						$\hat{\beta}_{F,j}$					
	j = 0	j = 1	j = 2	j = 3	j = 4	j = 5	j = 0	j = 1	j = 2	j = 3	j = 4	j = 5
AT	0.567***	0.461***	0.351	0.319	0.327	0.302	0.0583***	0.0659*	0.0529	0.0365	0.0343	0.0292
	[0.103]	[0.163]	[0.225]	[0.249]	[0.293]	[0.299]	[0.0200]	[0.0360]	[0.0591]	[0.0691]	[0.0968]	[0.0966]
BE	0.782***	0.741***	0.732***	0.742***	0.723***	0.712**	0.0369	-0.00627	-0.0356	-0.0330	-0.0335	-0.0350
	[0.113]	[0.168]	[0.212]	[0.246]	[0.255]	[0.286]	[0.0386]	[0.0721]	[0.0600]	[0.0762]	[0.0712]	[0.0783]
DE	0.782***	0.703***	0.677***	0.641***	0.640***	0.610***	0.0185	0.0484***	0.0741***	0.0865***	0.0898**	0.106**
	[0.0593]	[0.0521]	[0.0577]	[0.0486]	[0.0448]	[0.0502]	[0.0142]	[0.0167]	[0.0238]	[0.0325]	[0.0389]	[0.0466]
EL	0.759***	0.795***	0.716***	0.649***	0.550**	0.467	-0.00645	0.0119	0.0578	0.0973	0.0901	0.0856
	[0.113]	[0.146]	[0.202]	[0.242]	[0.266]	[0.290]	[0.0882]	[0.0877]	[0.0832]	[0.0886]	[0.0994]	[0.129]
ES	0.138**	0.0775*	0.0649	0.0603	0.0524	0.0459	0.0385	0.0318	0.0279	0.0130	0.000684	-0.0112
	[0.0582]	[0.0458]	[0.0475]	[0.0396]	[0.0387]	[0.0396]	[0.0305]	[0.0405]	[0.0492]	[0.0569]	[0.0568]	[0.0597]
FI	0.824***	0.638***	0.506***	0.499***	0.495***	0.469***	0.0904	0.102***	0.0462	0.0197	-0.00499	-0.0139
	[0.0486]	[0.0798]	[0.104]	[0.111]	[0.117]	[0.139]	[0.0588]	[0.0298]	[0.0474]	[0.0618]	[0.0598]	[0.0562]
FR	0.811***	0.706***	0.706***	0.690***	0.637***	0.635***	0.0684**	0.0978***	0.101***	0.105***	0.154***	0.178***
	[0.0944]	[0.0860]	[0.0938]	[0.106]	[0.128]	[0.144]	[0.0328]	[0.0277]	[0.0296]	[0.0336]	[0.0366]	[0.0437]
IT	0.550***	0.461***	0.437***	0.420***	0.413***	0.448***	0.0779***	0.110***	0.120***	0.120***	0.118***	0.110**
	[0.0298]	[0.0299]	[0.0365]	[0.0426]	[0.0520]	[0.0540]	[0.0288]	[0.0294]	[0.0309]	[0.0344]	[0.0396]	[0.0429]
NL	0.944***	0.977***	1.058***	0.997***	0.926***	0.823***	-0.0599	-0.0147	0.0427	0.0393	0.0383	0.0433
	[0.0725]	[0.0732]	[0.151]	[0.147]	[0.124]	[0.113]	[0.0844]	[0.109]	[0.108]	[0.109]	[0.112]	[0.118]
PT	0.295	0.157	0.0920	0.0433	0.0471	0.0666	0.0317	0.0562	0.0530	0.0800	0.0931	0.0943
	[0.183]	[0.139]	[0.109]	[0.116]	[0.114]	[0.155]	[0.0872]	[0.0837]	[0.101]	[0.113]	[0.0969]	[0.0973]
Observations	2790	2635	2480	2325	2170	2015	2787	2631	2476	2321	2167	2013

Note: This table reports the estimation of β_K and β_F using differentiated intervals, $j = 0 \ldots 5$. Standard errors are bootstrapped using 1000 iterations. * / ** / *** indicate 1% / 5% / 10% significance level.

Table B2: Estimation of the $\hat{\beta}_C$ and $\hat{\beta}_U$-coefficients

	$\hat{\beta}_{C,j}$						$\hat{\beta}_{U,j}$					
	j = 0	j = 1	j = 2	j = 3	j = 4	j = 5	j = 0	j = 1	j = 2	j = 3	j = 4	j = 5
AT	0.0666	0.0999	0.140*	0.172**	0.180*	0.202**	0.308***	0.374***	0.456***	0.472***	0.459***	0.467***
	[0.0557]	[0.0673]	[0.0822]	[0.0856]	[0.0979]	[0.101]	[0.0736]	[0.0946]	[0.123]	[0.129]	[0.146]	[0.151]
BE	0.0103	0.0233	0.0273	0.0113	0.0219	0.0296	0.171	0.242	0.276	0.280	0.288	0.293
	[0.0240]	[0.0150]	[0.0169]	[0.0210]	[0.0266]	[0.0347]	[0.139]	[0.183]	[0.215]	[0.231]	[0.239]	[0.248]
DE	-0.00788	-0.0127	-0.00656	0.00353	0.00428	0.00585	0.207***	0.262***	0.255***	0.269***	0.266***	0.278***
	[0.0107]	[0.0152]	[0.0147]	[0.0157]	[0.0168]	[0.0178]	[0.0537]	[0.0457]	[0.0394]	[0.0336]	[0.0344]	[0.0366]
EL	0.0381	0.0286	0.0433	0.0556	0.0924	0.123	0.198**	0.152	0.183	0.191	0.246	0.292
	[0.0400]	[0.0367]	[0.0447]	[0.0593]	[0.0710]	[0.0773]	[0.0957]	[0.116]	[0.140]	[0.170]	[0.208]	[0.241]
ES	0.140	0.0956	0.0791	0.0630	0.0688	0.0688	0.684***	0.795***	0.828***	0.864***	0.878***	0.897***
	[0.0879]	[0.0930]	[0.108]	[0.123]	[0.133]	[0.147]	[0.0962]	[0.106]	[0.111]	[0.123]	[0.121]	[0.140]
FI	0.0547	0.201***	0.468**	0.569**	0.582**	0.596**	0.0303	0.0588	-0.0201	-0.0876	-0.0727	-0.0506
	[0.0575]	[0.0722]	[0.220]	[0.281]	[0.282]	[0.295]	[0.0582]	[0.0874]	[0.139]	[0.206]	[0.206]	[0.259]
FR	0.0404*	0.0595**	0.0477*	0.0491	0.0542	0.0478	0.0802	0.137***	0.146**	0.156*	0.155*	0.139
	[0.0244]	[0.0245]	[0.0277]	[0.0322]	[0.0388]	[0.0417]	[0.0616]	[0.0509]	[0.0625]	[0.0813]	[0.0907]	[0.104]
IT	0.188***	0.149***	0.111	0.0965	0.0822	0.0551	0.184***	0.280***	0.332***	0.364***	0.387***	0.386***
	[0.0451]	[0.0563]	[0.0692]	[0.0817]	[0.0939]	[0.105]	[0.0289]	[0.0487]	[0.0700]	[0.0861]	[0.101]	[0.113]
NL	-0.0328	-0.0532	-0.0739	-0.0666	-0.0529	-0.0424	0.148	0.0913	-0.0266	0.0307	0.0888	0.176
	[0.0434]	[0.0422]	[0.0553]	[0.0606]	[0.0538]	[0.0511]	[0.0955]	[0.0669]	[0.0967]	[0.107]	[0.132]	[0.144]
PT	0.104**	0.125***	0.131**	0.145**	0.141**	0.149*	0.570***	0.662***	0.724***	0.732***	0.718***	0.690***
	[0.0513]	[0.0471]	[0.0584]	[0.0670]	[0.0687]	[0.0812]	[0.164]	[0.115]	[0.117]	[0.135]	[0.130]	[0.142]
Observations	2787	2631	2476	2321	2167	2013	2790	2635	2480	2325	2170	2015

Note: This table reports the estimation of β_C and β_U using differentiated intervals, $j = 0 \ldots 5$. Standard errors are bootstrapped using 1000 iterations. * / ** / *** indicate 1% / 5% / 10% significance level.

C Data sources

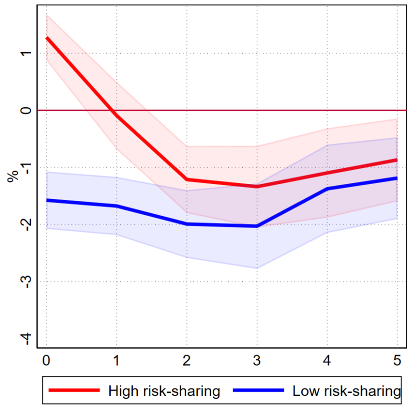
Table C1: Data sources

	Variable	Level	Note	Source
Risk-sharing	GDP	Regional	ln	Eurostat
	Primary income	Regional	ln	Eurostat
	Disposable income	Regional	ln	Eurostat
	Consumption	Regional	ln	Oxford Economics
Monetary policy	Short-term interest rate	Euro area	percent per annum	AWM database
	Shadow interest rate	Euro area	percent per annum	Lemke and Vladu (2017)
Control variables	Population	Regional	ln	Eurostat
	HICP	National	ln	Eurostat
	Stock market index	National	ln	OECD
	Government debt	National	% of GDP	Eurostat
	10y gov bond yield	National		ECB
	Structural primary balance	National	First-diff	AMECO
	GDP	Euro area	ln	Eurostat
	HICP	Euro area	ln	Eurostat

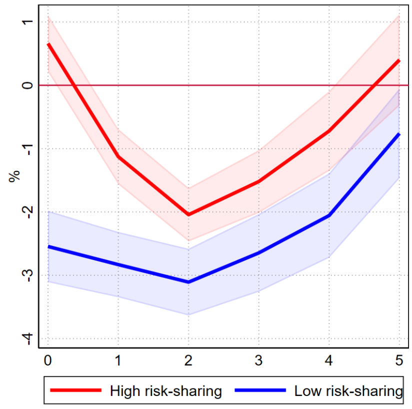
D Inequality and risk-sharing

Figure 5: Impact of monetary policy on output of poor vs rich regions when risk is shared through private risk-sharing channels

Factor market channel

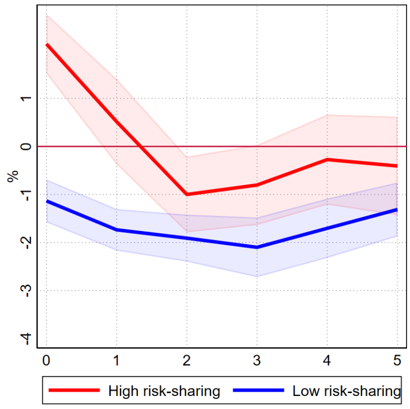


(a) Effect on poor regions (10th pct)

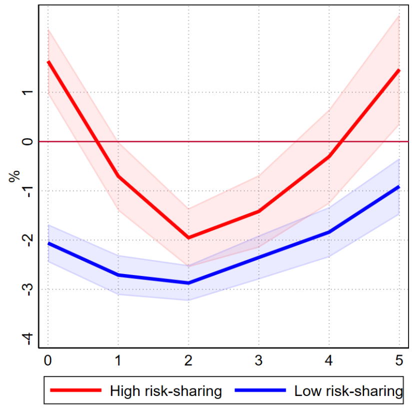


(b) Effect on rich regions (90th pct)

Credit market channel



(c) Effect on poor regions (10th pct)



(d) Effect on rich regions (90th pct)

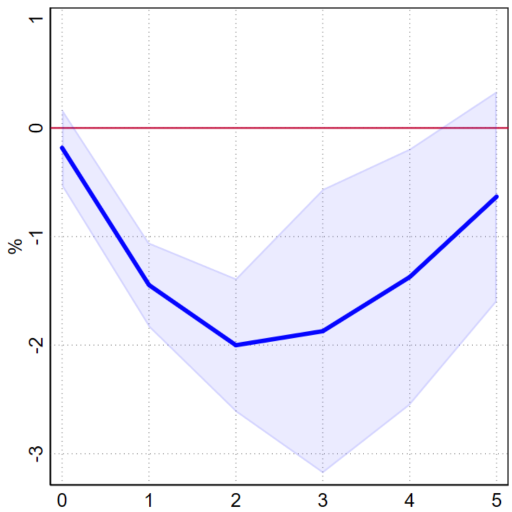
Note: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. Red (blue) lines depict the estimates for the upper (lower) deciles of $\hat{\beta}^{c}_{K,h}$ and $\hat{\beta}^{c}_{C,h}$. Standard errors are clustered at the time and regional-level and are boostrapped using 1000 iterations.

E Robustness checks

In this section, we report the impulse response functions of the robustness checks described in Section 5.

E.1 Excluding the largest regions from the regional sample

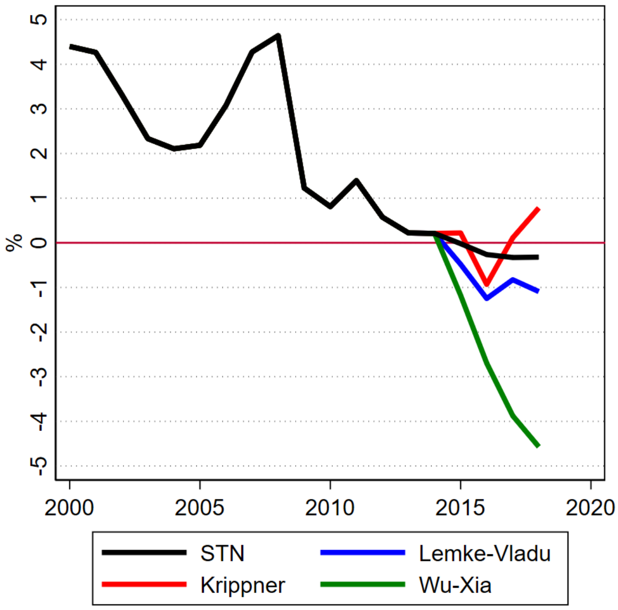
Figure 6: Impact of monetary policy on regional aggregates when excluding the largest regions



Note: 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. The 20 largest regions are excluded for each year.

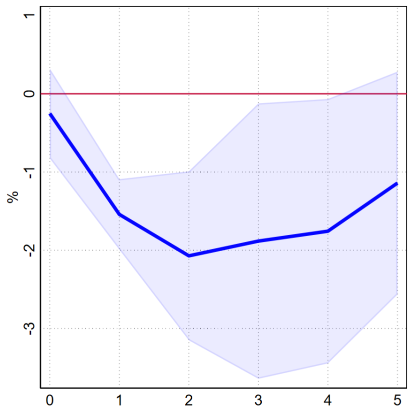
E.2 Alternatives for the shadow interest rate

Figure 7: Shadow rates for the Euro area

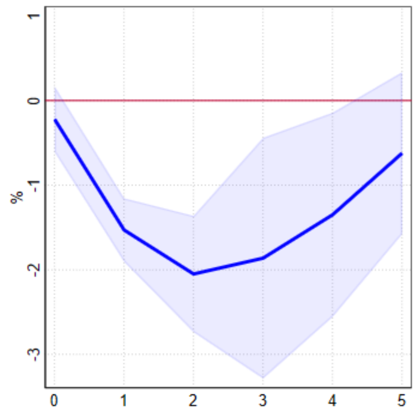


Note: The short-term interest rate (STN) is extended by adding the cumulative changes of the shadow rates developed by Lemke and Vladu (2017), Krippner (2015) and Wu and Xia (2017)

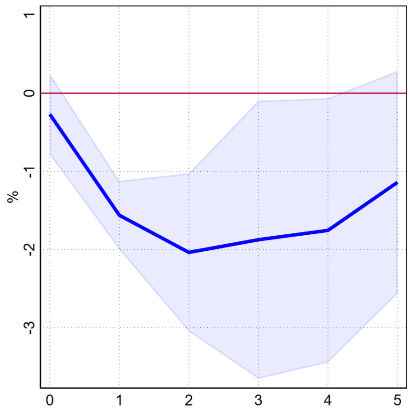
Figure 8: Impact of monetary policy on regional output using different monetary policy rates



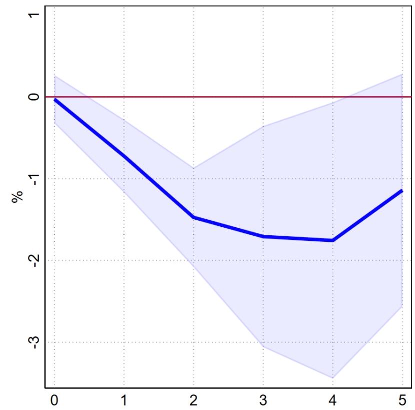
(a) Short-term interest rate



(b) Lemke and Vladu (2017)



(c) Krippner (2015)

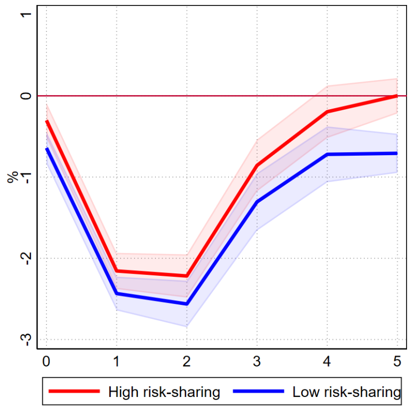


(d) Wu and Xia (2017)

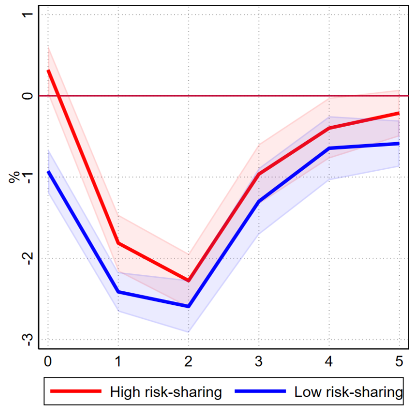
Note: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. The short-term interest rate (STN) is extended by adding the cumulative changes of the shadow rates developed by Lemke and Vladu (2017), Krippner (2015) and Wu and Xia (2017)

E.3 Adding oil prices and the real effective exchange rate

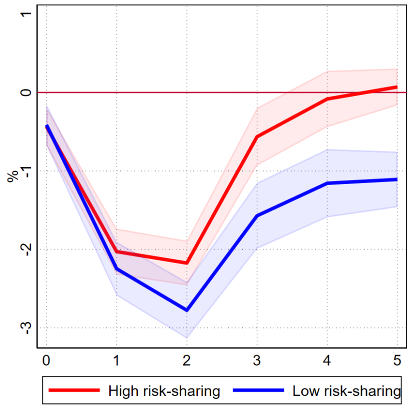
Figure 9: Impact of monetary policy when risk is shared with oil prices and REER



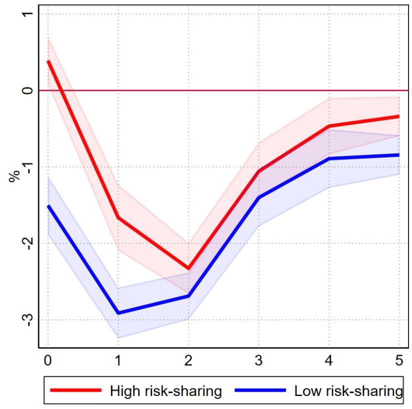
(a) All channels



(b) Factor market channel



(c) Fiscal channel

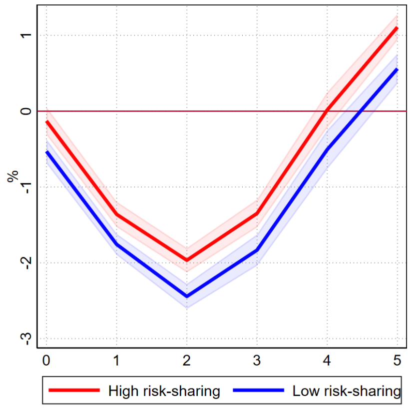


(d) Credit market channel

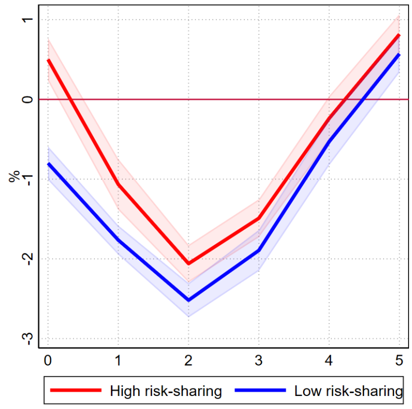
Note: Vertical axis refers to impact of 100 basis point rate hike on regional GDP (in %). Horizontal axis refers to horizon of IRF (in years). Solid lines denote point estimates and shaded areas denote 90% confidence bands. Red (blue) lines depict the estimates for the upper (lower) quartiles of $\beta^c_{S,h}$, $\beta^c_{K,h}$, $\beta^c_{F,h}$ or $\beta^c_{C,h}$. The Driscoll and Kraay (1998) standard errors are boostrapped using 1000 iterations.

E.4 Alternative control variables

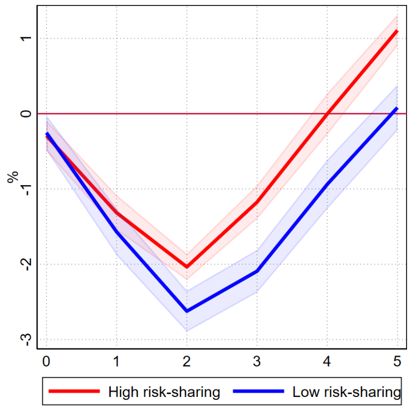
Figure 10: Impact of monetary policy when risk is shared with alternative control variables



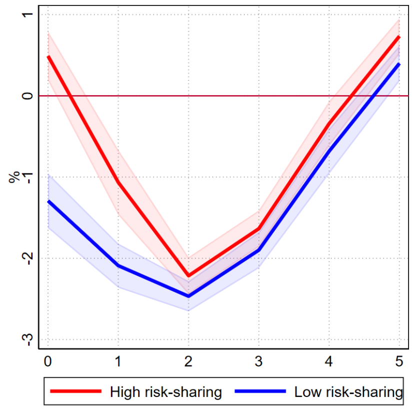
(a) All channels



(b) Factor market channel



(c) Fiscal channel



(d) Credit market channel

Note: 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. Red (blue) lines depict the estimates for the upper (lower) quartiles of $\beta^c_{S,h}$, $\beta^c_{K,h}$, $\beta^c_{F,h}$ or $\beta^c_{C,h}$. The Driscoll and Kraay (1998) standard errors are boostrapped using 1000 iterations.

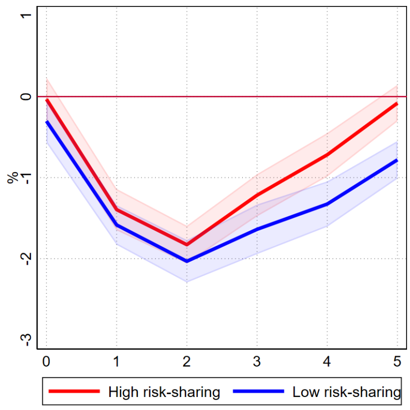
E.5 Forward-looking variables

We construct the forward-looking indicators for GDP and HICP from the quarterly ECB staff macroeconomic projections (see Hauptmeier et al. (2020) for further details). 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. We compute the average across the four projection vintages of each year for two years ahead. For most of the sample period, the horizon of the March, June, and September projection vintages of year t stretched until year $t+2$. To do so, we first calculate a forward-looking inflation variable as:

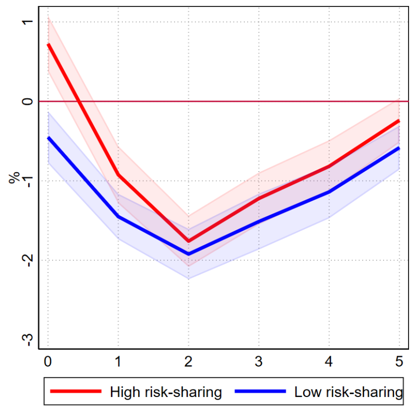
$$\pi^e_{t+2|t} = \frac{1}{4}(\pi^e_{t+2|DEC,t-1} + \pi^e_{t+2|MAR,t} + \pi^e_{t+2|JUN,t} + \pi^e_{t+2|SEP,t})$$

where $\pi^e_{t+2|t}$ is the expected inflation rate for year $t+1$ entering the central bank information set in year t, $\pi^e_{t+2|X,t-1}$ is the expected annual inflation rate in year $t+2$ according to the ECB macroeconomic projections in month X. The forward-looking variable for the rate of real economic growth is calculated in analogous fashion. We then compute the expected level of HICP using the respective inflation rate in year t and 2015 as the reference year (2015 = 100). Similarly, we calculate the expected level of GDP using the projected level of GDP in 2015. Consistent with the baseline specification, expected HICP and GDP are expressed in 100 times their log-levels.

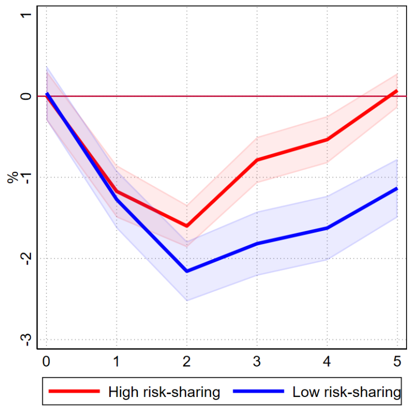
Figure 11: Impact of monetary policy when risk is shared with alternative control variables



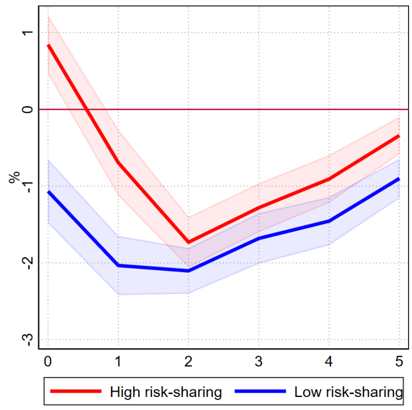
(a) All channels



(b) Factor market channel



(c) Fiscal channel



(d) Credit market channel

Note: 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. Red (blue) lines depict the estimates for the upper (lower) quartiles of $\beta^c_{S,h}$, $\beta^c_{K,h}$, $\beta^c_{F,h}$ or $\beta^c_{C,h}$. The Driscoll and Kraay (1998) standard errors are boostrapped using 1000 iterations.