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The Effectiveness of European Monetary Policy amid Global Supply Chain Pressures

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Abstract

This study employs nonlinear adaptations of local projection methods to explore how global supply chain conditions influence the transmission of monetary policy in the Euro Area and its four largest economies: Germany, France, Italy, and Spain. Finally, it leverages Denmarks unique monetary framework to exploit an exogenous source of variation, grounded in the trilemma of international finance, to strengthen the analysis of inflation dynamics. The findings reveal that heightened supply chain pressures do not consistently amplify the standard effects of monetary policy shocks on macroeconomic outcomes. While industrial production and inflation show a mildly increased sensitivity to monetary interventions under supply chain stress, no systematic state-dependent patterns emerge for retail trade and unemployment. To interpret these results, I empirically assess how the cost of external financing reacts to monetary policy across different supply chain conditions, testing the theory that an intensified credit channel underlies the observed amplification. Overall, the credit channels role appears secondary in the Euro Area. Two key insights emerge from a monetary policy perspective. First, advanced economies with credible monetary frameworks better anchor inflation expectations and stabilize inflation amid global supply chain disruptions, whereas emerging and low-income economies - characterized by weaker monetary frameworks - face greater challenges, increasing the risk of inflation expectations becoming de-anchored and diminishing monetary policy effectiveness. Second, the ambiguous effects on real economic variables limit the ability to quantify the economic costs of inflation stabilization under global supply chain disruptions. Further research in this direction would be highly valuable.

Keywords: Monetary policy; Supply chain disruptions; Credit channel; Non-linear local projections

JEL Classification: C32; E23; E30; E31; E32; E52

I would like to thank Professor Massimiliano Marcellino for the guidance provided. I also wish to thank my family and friends for their constant support throughout these years.

Contents

1 Introduction

It is well established that disruption of global supply chains[1] following the initial COVID-19 outbreaks has contributed significantly to increased inflation. However, there is limited understanding of how the functioning of global supply chains influences the transmission of monetary policy. As inflation increasingly becomes a global phenomenon, driven largely by factors beyond the control of domestic central banks, some may hastily conclude that central banks have inherently lost their ability to regulate inflation effectively. Motivated by such bold claims, this study seeks to examine how the effectiveness of monetary policy is shaped by the underlying conditions of global supply chains. Specifically, my objective is to determine whether stabilizing inflation amid global supply chain pressures becomes easier, remains equally challenging, or becomes more difficult.

This study is related to several areas of the literature. First, it connects to the literature on monetary policy and supply chains (Andriantomanga et al. (2023), Ascari et al. (2024), Bai et al. (2024), Carvalho et al. (2021), Ghassibe (2021), Nakamura and Steinsson (2010), Ozdagli and Weber (2017), Pasten et al. (2020), Wei and Xie (2020)). Second, it relates to the literature investigating the links between monetary policy and financial frictions (Bean et al. (2002), Bernanke and Gertler (1995), Bernanke (1999), Gertler and Karadi (2011), Gertler and Karadi (2015)). Third, this study contributes to the empirical macroeconomics literature on nonlinearities and state-dependent effects (Auerbach and Gorodnichenko (2012), Garcia and Schaller (2002), Peersman and Smets (2002), Peersman and Smets (2005), Passos and de Melo Modenesi (2021), Ramey and Zubairy (2014), Tenreyro and Thwaites (2016), Weise (1999)). Finally, and most importantly, it is also closely related to empirical research that examines the impact of global factors on monetary policy (Boivin and Giannoni (2008), Georgiadis and Mehl (2016)), with a particular focus on supply chain disruptions. To my knowledge, only two studies - Laumer and Schaffer (2025) and Hernández et al. (2024) - have analyzed the effectiveness of monetary policy amid global supply chain disruptions, focusing on the US and Mexico, respectively. This study contributes to this narrower strand of literature by adapting their analyses using data from the Euro Area, where this research question has not been previously explored, despite the strong participation of Europe in global supply chain networks. Furthermore, econometric analyses will be conducted in the four most populous European countries - Germany, France, Italy, and Spain - to test the robustness of the results and uncover potential heterogeneity and additional insights. Finally, the study offers a further contribution by taking advantage of Denmark's unique monetary framework to strengthen and corroborate the findings.

To address my research question, I implemented three different econometric models based on nonlinear adaptations of Jordà (2005) local projection framework.

The first analysis builds on Laumer and Schaffer (2025) and estimates state-dependent impulse response functions (IRFs) for industrial production, Consumer Price Index (CPI)[2], unemployment, and retail trade under both average and tight supply chain conditions, using data from the Euro Area

[1] In this study, I will use the terms "global value chains" and "global supply chains" interchangeably.

[2] I will use interchangeably CPI and HICP.

and individual countries. Unlike Laumer and Schaffer (2025), who highlight a strong amplification mechanism, whereby elevated supply chain pressures intensify the standard effects of monetary policy shocks through an intensification of the credit channel, my first analysis presents mixed evidence. Although a modest amplification effect appears for industrial production and a weaker one appears for the CPI, no clear pattern is observed for retail trade and unemployment. In these cases, IRFs appear volatile, neutral, or even economically insignificant in certain cases without systematic state-dependent differences.

Given the European Central Bank's primary mandate of maintaining price stability, and motivated by the weak statistical significance of the CPI result of the first analysis, the second and third analyses focus on estimating the accumulated impulse response functions of CPI inflation using instrumental variable techniques. Specifically, the second analysis builds on the methodology of Hernández et al. (2024), applying it to data from the Euro Area and individual countries. The third analysis draws on the approaches of Jordà et al. (2020), Hernández et al. (2024), and Di Giovanni et al. (2009) to conduct a similar investigation for Denmark to corroborate and strengthen the results. This third analysis leverages Denmarks fixed exchange rate regime and free capital mobility to exploit an exogenous source of variation, grounded in the trilemma of international finance. In line with Laumer and Schaffer (2025), the second and third analyses document a modest amplification mechanism for CPI, contradicting the findings of Hernández et al. (2024), who argued that monetary policy is less effective in stabilizing inflation amid global supply chain disruptions, or in other words that a given monetary policy shock has a weaker deflationary effect when supply chains are under stress. Overall, my results reveal minimal heterogeneity across the analyzed samples. A modest amplification effect is observed in the Euro Area, France, Spain, and Italy, a weaker one in Denmark, and no state-dependent difference in Germany.

Finally, I seek to interpret my results by testing the hypothesis proposed by Laumer and Schaffer (2025) using data from the Euro Area, namely that the amplification effect is driven by an intensified role of the credit channel. An empirical test is developed to verify the validity of this hypothesis. This last analysis does not invalidate the intensification of the credit channel mechanism, yet it suggests that its role appears to be secondary in the Euro Area and its constituent countries.

Despite the magnitude of the amplification effect, focusing solely on the inflation response, my analysis is more closely aligned with Laumer and Schaffer (2025) and Bai et al. (2024), confirming that the sensitivity of inflation to monetary policy increases when the global supply chain is under stress. At the same time, my findings diverge from those of Hernández et al. (2024) and Andriantomanga et al. (2023), who, examining similar questions in emerging and low-income economies, report that the influence of monetary policy on inflation is significantly dampened during supply chain disruptions. This attenuation is attributed to the important role of second-round effects on both non-tradable inflation and inflation expectations in these economies. Putting everything together, evidence suggests heterogeneous effects in the interaction between inflation response and global supply chain disruptions, yet a first pattern starts to emerge. On the one hand, advanced economies with credible monetary policy frameworks appear to be more capable of anchoring inflation expectations and more effectively controlling inflation. On the other hand, emerging and low-income economies struggle to stabilize

inflation due to weaker, less credible monetary frameworks, which increase the risk of inflation expectations de-anchoring.

When shifting the focus to the impact of monetary policy on real variables, no clear pattern emerges. Although a modest amplification mechanism is observed for industrial production, the response of the unemployment rate and retail trade remains ambiguous or statistically insignificant, with no systematic state-dependent effect identified. Understanding the effects of monetary policy amid global supply chain disruptions on these real variables is crucial, as it would help assess the economic costs of stabilizing inflation in similar scenarios. If the amplification mechanism proposed by Laumer and Schaffer (2025) is confirmed, reducing inflation would come at the cost of lower industrial production and retail trade, alongside higher unemployment. Conversely, if global supply chain stress renders output more inelastic, as suggested by Bai et al. (2024) - echoing the seminal findings of Keynes (1940) - such economic and social costs may not materialize. My analysis focuses primarily on inflation dynamics, aligning with the European Central Banks primary mandate to maintain price stability, and does not provide a definitive answer to this question. Further research in this direction would be highly valuable.

The remainder of this study is organized as follows. Section 1 presents the motivating evidence and highlights the relevance of the research question within the Euro Area context; Section 2 describes the data employed; Section 3 outlines the econometric methodology; Section 4 presents and discusses the empirical results; Section 5 tests Laumer and Schaffer (2025)'s hypothesis regarding the intensification of the credit channel; and Section 6 concludes the study.

1.1 Motivating Evidence

Before diving into the econometric section of this study, I would like to provide evidence that supports and motivates my research question.

1.1.1 Is Inflation Becoming an Increasingly Global Phenomenon?

Since the 1990s, inflation rates in various countries have shown a growing tendency to align, potentially due to several factors. These include the adoption of comparable and credible monetary policies worldwide and the influence of global commodity price fluctuations. International input-output connections have also played a relevant role in amplifying the effects of foreign cost shocks, thereby contributing to the convergence of inflation trends. Auer et al. (2017) note that cross-border trade in intermediate goods and services is the main channel through which global economic slack influences domestic CPI inflation. Similarly, Auer et al. (2019) document that the cross-border propagation of sectoral cost shocks through input-output linkages contributes substantially to synchronizing producer price inflation (PPI) across countries. Dexter et al. (2005) argue that globalization is responsible for the apparent weakening of the relationship between excess demand and inflation. They also highlight that international trade exerts a distinct influence on inflation and plays a critical role in identifying the Phillips curve relationship between unemployment and inflation. Similarly, Forbes (2019) emphasizes the need for inflation models to account more explicitly and comprehensively for changes in the global

economy, with key parameters adapting over time. Finally, disruption of global supply chains after the initial outbreak of COVID-19 was a major factor driving the surge in inflation (Bańbura et al. (2023), Finck et al. (2024), Liu and Nguyen (2023), Carrière-Swallow et al. (2023), Benigno (2022), Ascari et al. (2024), Bai et al. (2024), Gordon and Clark (2023), Finck and Tillmann (2022), LaBelle and Santacreu (2022), Di Giovanni et al. (2022), Andriantomanga et al. (2023)), providing further evidence of inflation's increasingly globalized nature.

These aspects can influence the trade-offs central banks face while managing inflation. Indeed, the monetary policy implications of global factors that play a greater role in shaping domestic inflation are profound, as these factors are beyond the control of individual central banks. These implications require careful consideration, as different responses can lead to different policy outcomes. Building on these observations - especially the substantial role of global supply chain disruptions in the recent inflation surge - this study aims to examine the interplay between global supply chains, a key channel connecting national economies, and the monetary policy transmission mechanism in the Euro Area.

1.1.2 Global Value Chains and the Euro Area

In recent decades, production processes have undergone significant changes, driven by reduced transportation costs and fewer barriers to international trade. As a result, production stages that once took place within a single country are now dispersed globally. In essence, many firms now source intermediate inputs from locations where production is most efficient, transforming them into goods or services that can cross borders multiple times before reaching their final consumers (Gunnella et al. (2019)). The growth and success of global value chains have become a defining feature of the recent era of globalization. However, interconnected supply chains come with a significant drawback: the dense network of global sourcing leaves individual countries highly vulnerable to disruptions of global value chains. This weakness became starkly evident during the COVID-19 pandemic when even minor disruptions in production and logistics morphed into substantial macroeconomic shocks (Finck and Tillmann (2022)). In addition, disruptions to global supply chains, once relatively rare before the COVID-19 pandemic, have become increasingly frequent in many countries (Ascari et al. (2024)). The ongoing discussions around deglobalization, reshoring, prioritizing robustness over efficiency, environmental challenges, trade route disruptions, and rising geopolitical tensions suggest that global supply chain pressures are likely to remain volatile in the foreseeable future (Laumer and Schaffer (2025)). Furthermore, the recent election of Trump, who has promised significant tariff increases, further amplifies the uncertainty surrounding the current geopolitical and macroeconomic landscape, with potentially substantial implications for the future dynamics of global supply chains. This highlights the urgent need to better understand the macroeconomic consequences of supply chain disruptions and the new challenges facing monetary policymakers.

In addition, disruptions in global value chains have a particularly pronounced impact on the Euro Area due to its strong dependence on international supply networks. Figure 1 illustrates the evolution of the average share of GVC-related manufacturing output in the Euro Area from 2007 to 2022, compared to the same metric for several major non-European economies (Brazil, Canada, China, India,

Japan, Mexico, Russia and the USA). While Euro Area countries exhibit a clear upward trend, no comparable pattern emerges among the non-European economies. Additionally, a stark difference in levels is evident, with GVC-related manufacturing output in the Euro Area consistently at least twice as high as in the non-European countries throughout the entire period.

Reinforcing the importance of this topic for the Euro Area, Figure 2 provides more detailed information at the country level. As shown, in many countries of the Euro Area, more 40% of the output of the manufacturing sector - either directly or indirectly - crosses multiple borders. In contrast, non-European countries show significantly lower percentages, highlighting the stronger integration of the Euro Area into global value chains.

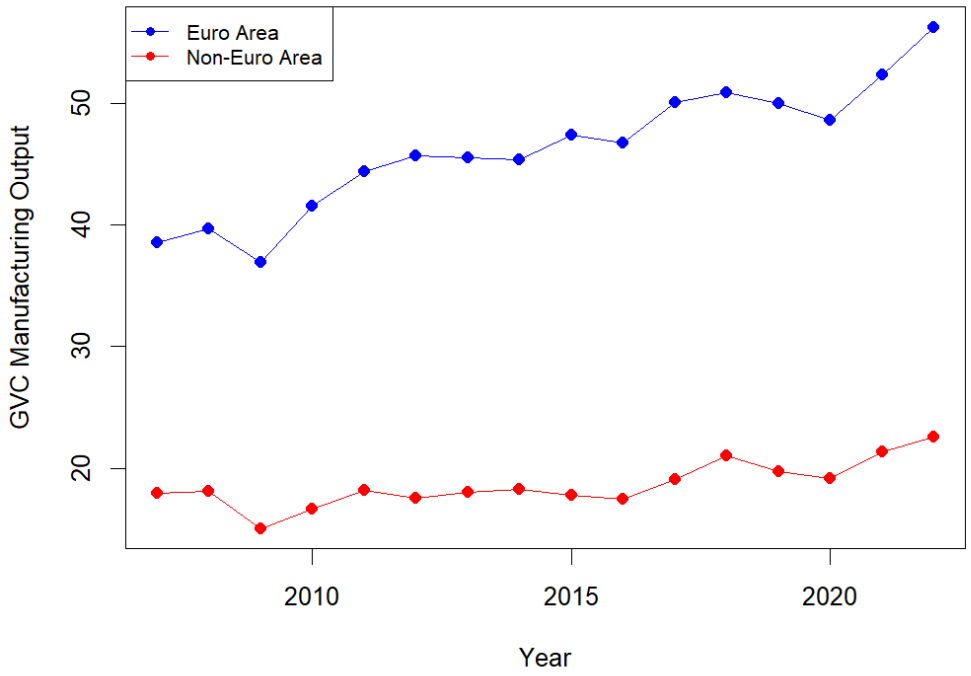


Figure 1: Evolution of GVC Manufacturing Output

Consequently, global supply chain disruptions can become a major driver of business cycle fluctuations, and participation in global value chains has substantial macroeconomic implications for the economy of the euro area (Finck and Tillmann (2022)). This further underscores the need to adapt traditional macroeconomic analysis and forecasting approaches for the Euro Area to better understand the macroeconomic consequences of global value chain disruptions.

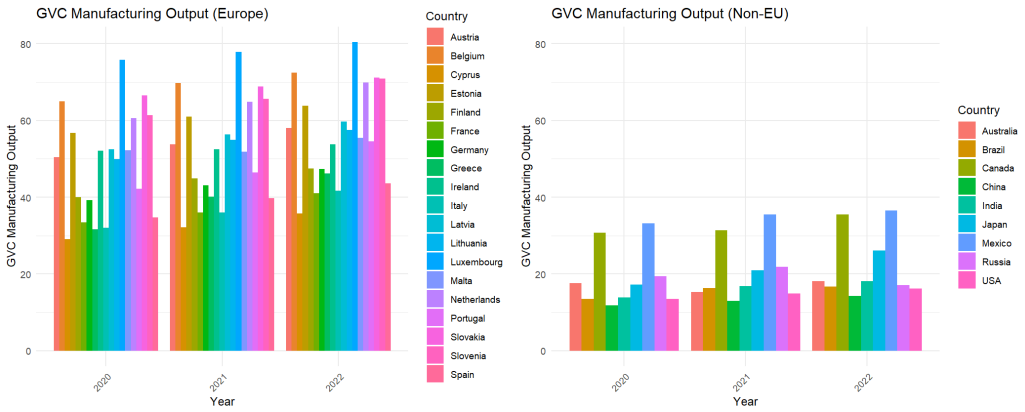


Figure 2: Evolution of GSCPI and Inflation Rate

Benigno (2022) developed an index, the Global Supply Chain Pressure Index (GSCPI), to capture the status of global supply chains. The index is normalized such that a zero indicates that the index is at its average value with positive (negative) values representing how many standard deviations the index is above (below) this average value.

The GSCPI offers an approximate measure of potential imbalances between demand and supply that arise specifically from supply disruptions. The authors developed this index by isolating and combining supply-side factors from country-specific supply chain data (China, Japan, Korea, Taiwan, UK, US and the Euro Area) and global transportation cost metrics. Figure 3 illustrates the evolution of the GSCPI alongside the European inflation rate (CPI).

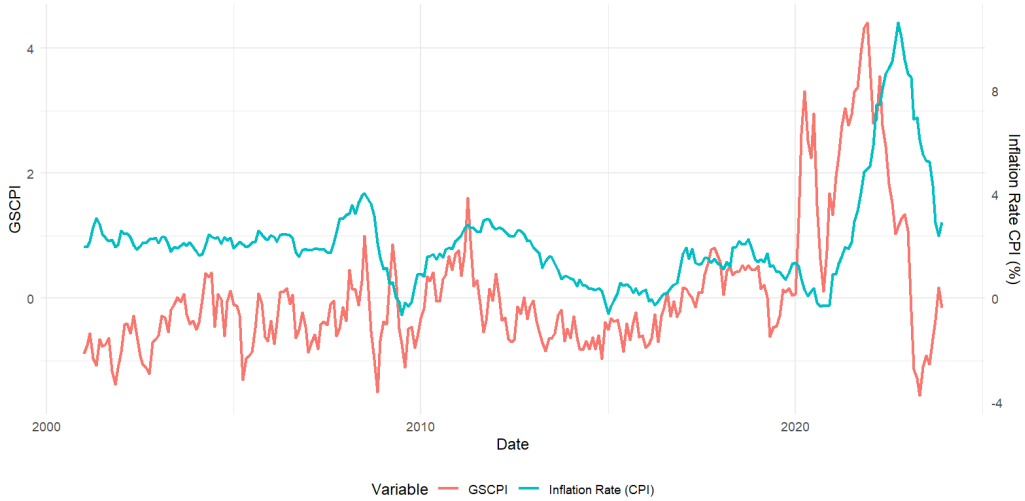


Figure 3: Evolution of GSCPI and Inflation Rate (CPI)

The GSCPI experienced an increase during the Global Financial Crisis, although Benigno (2022) caution that not all demand-related effects may have been fully excluded from the index during that period. The index increased significantly in 2011 due to the Tohoku earthquake and subsequent tsunami, which affected the Fukushima nuclear reactor. It also remained elevated from 2017 to 2019 as

a result of flooding in Thailand and the US-China trade war, both of which caused global supply chain disruptions. However, the most severe disruptions were driven by the recent COVID-19 pandemic, during which the index reached its highest recorded levels.

As shown in Figure 3, there is a strong positive correlation between the GSCPI and price dynamics and, as already mentioned, numerous studies confirm that global supply disruptions have significantly contributed to the recent surge in inflation. However, less is known about the impact of global supply chains on the transmission of monetary policy. To my knowledge, only Hernández et al. (2024) and Laumer and Schaffer (2025) have empirically examined how the functioning of global supply chains affects the transmission of monetary policy, specifically in Mexico and the United States. My objective is to conduct a similar analysis using European data, where this issue is particularly relevant and has not yet been explored.

1.1.3 Related Literature

I now provide additional evidence from the related literature, addressing similar research questions from both theoretical and empirical perspectives.

Ascari et al. (2024) examine how global supply chain pressures influence inflation dynamics in the Euro Area and explore the associated monetary policy implications. Their analysis indicates that global supply chain shocks were the primary driver of the surge in Euro Area inflation during the first half of 2022. They also highlight the persistent and hump-shaped nature of these shocks' effects on inflation. Notably, supply chain disruptions can lead to second-round effects, further elevating aggregate prices and sustaining inflationary pressures even after the initial impact of the shocks has dissipated. To investigate these dynamics, the authors employ a two-country New Keynesian model, focusing on the relationship between global supply chain pressures and optimal monetary policy. Their findings reveal a non-linear connection between GVC participation and the optimal monetary policy response to supply-induced inflation. At lower levels of GVC participation, supply chain pressures result in higher domestic inflation and output, necessitating a contractionary policy stance. Conversely, at higher GVC participation levels, supply chain pressures lead to rising inflation but declining output, complicating the trade-off between stabilizing inflation and output. In this case, a less aggressive monetary tightening is warranted to address these challenges effectively.

Andriantomanga et al. (2023) investigate the impact of supply chain pressures on headline, food, and tradable inflation in a panel of 29 sub-Saharan African countries between 2020 and 2022. Their findings reveal that supply chain disruptions significantly influenced these inflation components. The authors emphasize the importance of monetary policy responses to global value chain shocks, arguing that central banks can more effectively stabilize inflation and output by closely monitoring global value chains and preemptively adjusting their policy stance before the disruptions fully propagate through all inflation components. Proactive measures help mitigate second-round effects on non-tradable inflation and inflation expectations. These second-round effects are particularly significant in the region due to the high proportion of food and tradable goods in consumption baskets and the limited credibility of central banks, which increases the risk that inflation expectations become unanchored.

Bai et al. (2024) argue that supply chain shocks were the primary drivers of inflation in the United States during 2021, while traditional demand and supply factors played a more significant role from 2022 onward. Their analysis demonstrates that monetary policy becomes more effective at stabilizing inflation during periods of supply chain disruptions compared to normal conditions. Specifically, their model-based findings suggest that such disruptions heighten the sensitivity of prices to changes in demand while rendering output largely inelastic. In this context, contractionary monetary policy has a greater capacity to stabilize inflation with a reduced impact on output. They empirically validate these theoretical predictions, emphasizing that supply chain disruptions warrant a more aggressive tightening of monetary policy. Their conclusions closely align with Keynes (1940) seminal analysis, which posited that when output is constrained, policymakers can sharply reduce aggregate demand to combat inflation with minimal concern for its effects on production.

Wei and Xie (2020) explore the effects of global supply chains on optimal monetary policy within a small open economy New Keynesian framework that incorporates multiple stages of production. Their analysis reveals that as an economy becomes increasingly open, the optimal emphasis placed on upstream inflation grows relative to the emphasis on final-stage inflation. Furthermore, they demonstrate that focusing on PPI inflation results in a smaller welfare loss compared to targeting CPI inflation exclusively. In particular, as the production chain extends, the ideal weight assigned to PPI inflation in a policy rule that accounts for both PPI and CPI inflation also increases.

A growing body of research highlights the importance of production networks in shaping the effects of monetary policy. For example, Pasten et al. (2020), Nakamura and Steinsson (2010), and Carvalho et al. (2021) explore how production networks amplify strategic complementarities in price-setting, thereby increasing short-run money non-neutrality. In support of this perspective, Ghassibe (2021) provides empirical evidence showing that the amplification effects of the input-output link accounts for at least 30% of the overall impact of monetary shocks on aggregate consumption. Similarly, Ozdagli and Weber (2017) find that indirect production network links are responsible for at least half of the observed effect of monetary shocks on stock returns.

Research exploring the influence of global forces on the effectiveness of monetary policy remains relatively sparse. Boivin and Giannoni (2008) find limited evidence of global forces altering the transmission mechanism of monetary policy. Their point estimates suggest that the increasing prominence of global forces may have contributed to reducing some persistence in policy responses. However, they conclude that if global forces have impacted the monetary transmission mechanism, this is a relatively recent phenomenon. Similarly, Georgiadis and Mehl (2016) show that financial globalization has not significantly affected the effectiveness of monetary policy in the euro area since the late 1990s. In contrast, financial globalization has enhanced monetary policy effectiveness in most non-euro area advanced and emerging market economies.

As noted in Ascari et al. (2024), the impact of global supply chain disruptions on the monetary policy transmission mechanism, which is the central focus of this study, has not been adequately studied. Bai et al. (2024) touch on the effectiveness of contractionary monetary policy in the middle of supply chain disruptions, but their main focus is not on the transmission mechanism itself. In contrast, Laumer and Schaffer (2025), using a nonlinear adaptation of the local projection framework, find that height-

ened supply chain pressures amplify the standard effects of monetary policy shocks on macroeconomic aggregates. According to their analysis, this amplification occurs because credit costs become more sensitive to monetary policy shocks when global supply chains are under strain. Hernández et al. (2024) take a different approach by studying the impact of global supply chain disruptions on the ability of Mexico's central bank, representing an emerging market economy, to stabilize inflation. Using a non-linear local projection framework with monetary policy shocks instrumented by federal funds rate shocks, they find that during periods of supply chain stress, the impact of monetary policy on inflation over a one-year horizon is significantly weakened compared to non-stress scenarios. They attribute this reduced efficacy to the slow adjustment of inflation expectations in high-stress regimes. These findings contrast sharply with those of Laumer and Schaffer (2025) and, to some extent, align with the conclusions of Andriantomanga et al. (2023), who analyze 29 sub-Saharan African countries between 2020 and 2022. The latter emphasizes the importance of second-round effects on both non-tradable inflation and inflation expectations. They also highlight the reduced ability of central banks to stabilize inflation under supply chain distress, advocating for proactive monitoring of global supply conditions and, where necessary, preemptive adjustments to monetary policy to mitigate second-round effects.

2 Data

2.1 Data - First Analysis

The first econometric analysis, based on the work of Laumer and Schaffer (2025), includes the following macro variables sourced from the Euro indicators dashboard: [3]

- Harmonised index of consumer prices (HICP): [4] a set of consumer price indices calculated according to a harmonised approach and to definitions laid down in regulations.
- Industrial Production: the index of industrial production measures the evolution of the volume of production for industry excluding construction.
- Retail Trade: the index of the volume of retail trade measures the evolution of the total amount of sales, adjusted for price changes (deflated), i.e., the evolution of the total amount of goods sold.
- Unemployment Rate: defined as the percentage of the labor force constituted by persons aged 15 to 74 who: are without work; are available to start work within the next two weeks; and have actively sought employment at some time during the previous four weeks.

[3] `https://ec.europa.eu/eurostat/cache/dashboard/euro-indicators/`

[4] For the remainder of this study, I will use HICP and CPI interchangeably.

Furthermore, this initial analysis incorporates the monetary shock identified by Jarociski and Karadi (2020). In place of the excess bond premium (EBP) - for which no European equivalent exists, as noted by Jarociski and Karadi (2020) - I include the ICE BofA Euro High Yield Index Option-Adjusted Spread.[5]

As highlighted by Laumer and Schaffer (2025), the EBP represents the portion of the interest rate spread between a corporate bond index and a government bond of comparable maturity that is not attributable to default risk. As such, it captures the additional credit costs borne by private firms relative to government securities, purely due to financial frictions. Consequently, the EBP serves as a direct measure of the external finance premium for large corporations. My selected series should capture a similar economic mechanism, and for the sake of exposition, I will refer to it as the EBP. Finally, to account for global supply chain conditions, I use the Global Supply Chain Pressure Index (GSCPI) developed by Benigno (2022) outlined in the introductory section.

In the baseline analysis, following the approach of Laumer and Schaffer (2025) and considering data availability, the dataset spans from January 2001 to December 2019. For the COVID extension, I extend the sample period to include observations until December 2022.

2.2 Data - Second and Third Analysis

The choice of regressors for the second and third analyses follows the methodology established by Hernández et al. (2024). The first regression, designed to isolate the federal funds rate shock, includes:

- (Shadow) Federal funds rate developed by Wu and Xia (2016), [6]
- US industrial production, [7]
- Consumer Price Index, [8]
- The 1-year ahead inflation expectations, [9]
- The market 10-year yield on US Treasuries, [10]

[5] This series is freely available in the St. Louis FRED database under the ID BAMLHE00EHYIOAS. Jarociski and Karadi (2020) substitute the EBP with the BBB bond spread. However, since I could not find this specific series and following the suggestion of Marek Jarocinski, I opted for the ICE BofA Euro High Yield Index Option-Adjusted Spread. This spread is for bonds a notch more risky than BBB, but it very closely comoves with the BBB spread.

[6] `https://sites.google.com/view/jingcynthiawu/shadow-rates`

[7] `https://fred.stlouisfed.org/series/INDPRO`

[8] `https://fred.stlouisfed.org/series/CPIAUCSL`

[9] `https://fred.stlouisfed.org/series/EXPINF1YR`

[10] `https://fred.stlouisfed.org/series/DGS10#0`

In the second regression, the goal is to isolate exogenous variation in the monetary policy rate of the countries of interest, the Euro Area, Germany, France, Italy, Spain, and Denmark. To measure monetary policy, I use the EONIA from the Euro Indicators Dashboard.[11] The control variables include:

- The lag of the HICP, as defined in the previous subsection (log-difference).
- As a measure of domestic economic activity, accounting for demand-side price pressures, I include the Industrial Production Index, as previously defined, and the Construction Index[12], both sourced from the Euro Indicators Dashboard (log-difference).
- To control for domestic supply-side inflationary pressures, I use the Producer Price Index (PPI) from the St. Louis FRED database (log-difference).
- The exchange rate depreciation rate is included to account for both inflationary pressures in the tradable sector and interest rate pressures from capital flows. It is defined as 100 times the log-difference of the real effective exchange rate, sourced from the Euro Indicators Dashboard (log-difference).
- Following Hernández et al. (2024), I control for global variables that remain invariant between regimes. These include the log-difference of VIX[13], which measures financial market stress; the log-difference of WTI oil prices[14], which captures noncore inflationary pressures and potential secondary effects on the overall Consumer Price Index; and the log-difference of the G7 Industrial Production Index[15], which accounts for external demand pressures and supply constraints.

The third regression includes the same set of regressors, except that the dependent variable is defined as 100 times the logarithmic difference between HICP at time $t+h$ and HICP at time t. Due to data availability, the dataset spans from January 2001 to December 2021.

3 Econometric Section

To examine how the effectiveness of European monetary policy varies depending on the underlying state of global supply chain conditions, I will employ non-linear extensions of Jordà (2005) local projection methods. These methods are inspired by two recent studies, Laumer and Schaffer (2025) and

[11] The EONIA (Euro OverNight Index Average) is the effective overnight reference rate for the euro, computed as a weighted average of all overnight unsecured lending transactions in the interbank market, initiated within the euro area by the contributing panel banks. EONIA is computed by the European Central Bank.

[12] Production in construction reflects the output and activity of the construction sector.

[13] `https://fred.stlouisfed.org/series/VIXCLS\#0`

[14] `https://fred.stlouisfed.org/series/MCOILWTICO`

[15] `https://fred.stlouisfed.org/series/G7PRINTO01IXOBSAM`

Hernández et al. (2024), which explore similar research questions in different contexts. Furthermore, the third analysis is also specifically based on the work of Jordà et al. (2020) and Di Giovanni et al. (2009). Local projection methods are not only central to these articles, with the exception of Di Giovanni et al. (2009), but are also a prominent tool in the broader literature investigating nonlinear and state-dependent effects (see, for instance, Ramey and Zubairy (2014) and Passos and de Melo Modenesi (2021)). Local projections offer several advantages: they are robust to model misspecification and allow for experimentation with highly nonlinear and flexible specifications, making them particularly well suited to the objectives of this research.

The econometric analysis is organized into three subsections. The first subsection, following Laumer and Schaffer (2025), estimates impulse response functions for industrial production, CPI inflation, unemployment, and retail trade. Then, given the European Central Bank's primary mandate of maintaining price stability, and the low statistical significance of the CPI response in the initial analysis, the second and third subsections focus on estimating the accumulated impulse response function of CPI inflation using instrumental variable techniques. Specifically, the second subsection replicates the methodology of Hernández et al. (2024), itself inspired by Jordà et al. (2020). The third subsection extends this approach to Denmark, drawing on Jordà et al. (2020), Hernández et al. (2024), and Di Giovanni et al. (2009), to corroborate and strengthen the findings. Although the estimated equations in the third analysis are identical to those in the second, additional considerations support the instrumental variable strategy. In particular, Denmarks fixed exchange rate and free capital mobility provide a source of exogenous variation, grounded in the trilemma of international finance. The first two models will be applied to aggregate Euro Area data, as well as to data from the largest Euro Area economies - Germany, France, Italy, and Spain - to explore cross-country heterogeneity, check the robustness of the results, and derive additional insights.

3.1 First Analysis

The first analysis is based on the work of Laumer and Schaffer (2025).

3.1.1 Econometric Model

Using a non-linear adaptation of Jordà (2005) local projections, the alleged state dependence of monetary policy will be analyzed through an interaction term:

$$
\begin{aligned}
x_{t+h} &= c^h + \sum_{j=1}^{J} \alpha_j^h x_{t-j} + \sum_{j=0}^{J} \beta_j^h mp_{t-j} \\
&+ \psi_j^h GSCPI_{t-1} + \sum_{j=0}^{J} \delta_j^h GSCPI_{t-1} mp_{t-j} + \sum_{j=0}^{J} \Gamma_j^h Z_t + \epsilon_{t+h} \\
&for \quad h = 0, ..., H
\end{aligned}
$$

where c^h is a constant, x is the outcome variable of interest, Z is a vector of control variables,[16] GSCPI is the Global Supply Chain Pressure Index developed by Benigno (2022) described in the introductory section, mp is the high-frequency monetary policy shock identified by Jarociski and Karadi (2020), and ϵ is the residual. Note that, as outlined by Jordà (2005), ϵ_{t+h} represents a scalar moving average process of order h. Thus, the residuals are expected to exhibit autocorrelation. Although this does not compromise the consistency of the OLS estimates, it does affect the error variance matrix estimation, which in turn impacts the construction of confidence intervals. To account for autocorrelation in the residuals, I rely on the heteroskedasticity and autocorrelation-consistent covariance matrix estimators (HAC) based on Newey and West (1987). The only difference from Laumer and Schaffer (2025)'s original specification is that the control matrix Z includes only contemporaneous values, omitting lags.[17] The optimal number of lags is determined using the Bayesian Information Criterion (BIC), which suggests including two lags.[18]

My focus is on the set of coefficients $\{\beta_0^h\}_{h=0}^H$ and $\{\beta_0^h + \delta_0^h\}_{h=0}^H$ to analyze impulse responses. The first set of coefficients, $\{\beta_0^h\}_{h=0}^H$, represents the effect of a monetary policy shock at t on the variable of interest at $t+h$ when global supply chain pressures are at their average level - namely, under "normal" supply conditions. Instead, $\{\delta_0^h\}_{h=0}^H$ capture potential state dependencies by measuring the additional impact of a monetary policy shock when global supply chain pressures are elevated - specifically, one standard deviation above their average level.

In the standard specification of this model, I use the baseline monetary shocks identified by Jarociski and Karadi (2020). These shocks are identified through a combination of high-frequency identification (HFI), sign restrictions, and the assumptions that within a 30-minute window surrounding FOMC announcements, only two structural shocks, monetary and information shocks - systematically influence financial market surprises. The authors' idea is that central bank announcements simultaneously convey information about monetary policy actions and the central bank's assessment of the macroeconomic outlook. Jarociski and Karadi (2020) exploit the high-frequency comovement of interest rates and stock prices within a narrow time window around policy announcements to unravel the two distinct shocks. To achieve this, they impose the restriction that a monetary policy shock leads to higher interest rates and lower stock prices, whereas an information shock raises both variables. This approach leverages the fact that standard economic theory provides clear predictions about the

[16]The control variables include the three remaining macro variables for which the IRF is not being computed, along with my version of the EBP.

[17]I used this "different" model to replicate Laumer and Schaffer (2025)'s results and obtained nearly identical IRFs.

[18]Following the recommendation of Sebastian Laumer and Matthew Schaffer, I applied the BIC using the four macro variables as dependent variables, also shifting them forward in time as done in the Local Projections implementation. The most commonly suggested lag length across different model specifications was two. However, the results remain robust to variations in the number of lags included. I also applied this procedure to Laumer and Schaffer (2025)'s dataset using the Akaike Information Criterion (AIC) and obtained their suggested lag length, further validating my selection procedure.

direction of these comovements following a policy change. Neglecting the role of information shocks, they argue, can lead to biased conclusions about the non-neutrality of monetary policy.
The state variable, GSCPI, is lagged to prevent contemporaneous feedback between the state and the shocks. This approach is standard in the literature using state-dependent IRFs. For example, studies such as Auerbach and Gorodnichenko (2012), Passos and de Melo Modenesi (2021), Ramey and Zubairy (2014), and Hernández et al. (2024) also rely on lagged state variables.
All macroeconomic variables, except the unemployment rate, are expressed in logarithmic form. The monetary policy shock is normalized so that a unit change corresponds to a one-standard-difference change. Similarly, the GSCPI is standardized to have a mean of zero and a standard deviation of one over the sample period considered. Finally, $H = 36$ as in Laumer and Schaffer (2025), so I will investigate the effects of a monetary shock up to 3 years after its occurrence.

3.2 Second Analysis

The second analysis builds on the research of Hernández et al. (2024) and examines the trajectory of accumulated CPI inflation[19] in response to an exogenous monetary policy action, denoted as Δr_t. To estimate this response, I will employ a nonlinear version of Jordà (2005) local projection framework where I will control for a set of relevant macroeoconomic factors, x_t, also including lags of the dependent variable.[20] Neglecting alleged nonlinearities for now, the accumulated responses are derived from the estimated β_h coefficients obtained through a series of local projections:

$$y_{t+h} = \alpha_h + \Delta r_t \beta_h + x_t \gamma_h + v_{t+h}, \quad h = 1, ..., H \tag{1}$$

Again, as noted in Jordà (2005), v_{t+h} represents a scalar moving average process of order h, while α_h and γ_h are parameters. To account for the autocorrelation in the residuals, I again rely on the heteroskedasticity and autocorrelation-consistent (HAC) covariance matrix estimators based on Newey and West (1987).
Unlike the previous analysis, which used the monetary policy shocks identified by Jarociski and Karadi (2020), this approach measures monetary policy interventions as changes in the interbank interest rate (EONIA). However, these policy actions are not exogenous, as they respond to variations in CPI inflation and other macroeconomic factors. This endogeneity poses a challenge to the consistency of the standard OLS estimation for the coefficients β_h. To overcome this issue, an identification strategy using external instruments is applied.

3.2.1 Identification of Monetary Shocks

Following the approach outlined by Hernández et al. (2024), I identify an exogenous variation in the target interest rate through an instrumental variable method inspired by Jordà et al. (2020).

[19] $100(ln(CPI_{t+h}) - ln(CPI_t))$

[20] For a comprehensive description of the control variables, refer to the Data section.

Specifically, Hernández et al. (2024) adopts the idea from Jordà et al. (2020) of utilizing shocks [21] to the monetary policy rate of a base country as an instrument to isolate exogenous variations in the monetary policy rate of the country whose impulse response function is of interest. As argued later, if their empirical validity holds in the Mexican context, it should also be applicable to my sample.[22] As illustrated in the figure below, and consistent with the logic of uncovered interest rate parity, the monetary policy rate determined by the European Central Bank exhibits a strong correlation with that set by the Federal Reserve.

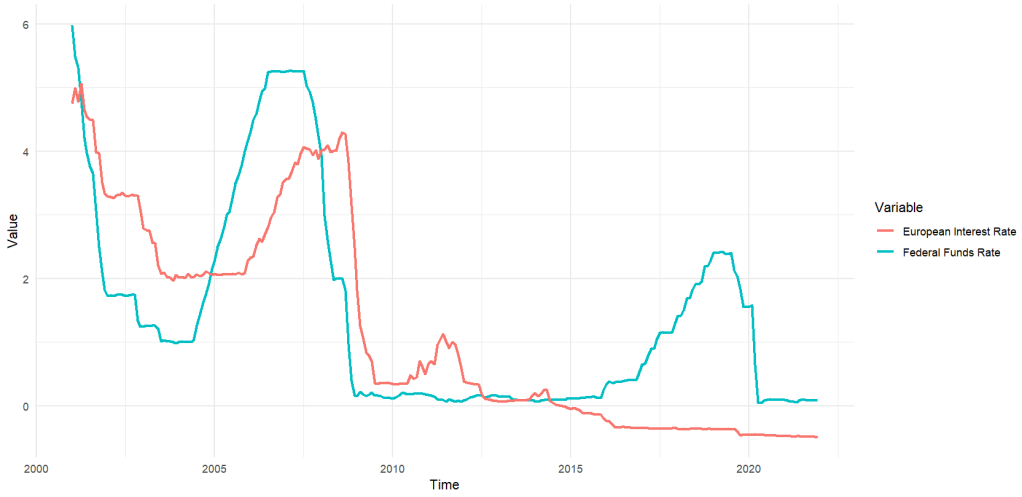


Figure 4: Evolution European and US Interest Rate

This motivates the following stepwise estimation strategy.
First, I estimate shocks to the (shadow) federal funds rate developed by Wu and Xia (2016) as the residuals of a regression of this rate on its primary determinants:

$$\Delta r_t^* = \alpha_h^* + x_t^* \gamma_h^* + \eta_t \tag{2}$$

where Δr_t^* is the first difference of the (shadow) federal funds rate, and x_t^* includes key macroeconomic variables: US industrial production, Consumer Price Index (both in log-difference), the 1-year ahead inflation expectations, and the market 10-year yield on US Treasuries. It is important to note that the estimated monetary policy shocks, the residuals $\widehat{\eta}_t$, are not correlated with US economic activity, US inflationary dynamics, and global financial conditions, as captured by the 10-year yield.
Thus, following Jordà et al. (2020), I use the fitted residuals $\widehat{\eta}_t$ from the previous regression as an

[21] By "shocks," I refer to the unpredictable components, or residuals, from a regression of the base country's monetary policy rate on its primary macroeconomic determinants.

[22] A more accurate replication of Jordà et al. (2020)'s empirical strategy necessitates an economy characterized by free capital mobility and a fixed exchange rate. For this reason, the third analysis estimates state-dependent impulse response functions (IRFs) in Denmark, an economy that meets these criteria. Still, this empirical strategy remains valid also in the European context as extensively argued later on.

instrument to identify the exogenous variation in Δr_t, the European Interest Rate (EONIA), in the following regression:

$$\Delta r_t = \delta_0 + \widehat{\eta}_t \psi + x_t \delta + u_t \tag{3}$$

from which I obtain the exogenous fitted values $\widehat{\Delta r_t}$. These are exogenous variations in the European Interest Rates that will finally be used in the local projections to approximate the IRFs:

$$y_{t+h} = \alpha_h + \widehat{\Delta r_t} \beta_h + x_t \gamma_h + v_{t+h}, \quad h = 1, ..., H \tag{4}$$

It is important to note that, for the sake of argument, the above expression does not incorporate the nonlinear specification required for the analysis. In addition, note that x_t includes a sets of control variables, extensively described in the Data section, also including 1 lag of the dependent variable, where the number of lags is selected applying the Bayesian Information Criterion (BIC). [23]

Following Hernández et al. (2024)'s arguments, I now argue why this is a valid instrument. Specifically, I focus on the relevance, contemporaneous exogeneity, and lead-lag exogeneity conditions outlined in Stock and Watson (2018) and Jordà (2023).

Regarding relevance, I find that $cor(\Delta r_t, \widehat{\eta}_t) = 0.15$ with a t-statistic of 2.34, indicating that the correlation is statistically different from zero.

With respect to contemporaneous exogeneity, one could argue that shocks to the federal funds rate are uncorrelated with macroeconomic conditions in the Euro Area, particularly with CPI inflation. Given the dominant role of the US economy, it is unlikely that European consumer inflation significantly influences the Federal Reserve's monetary policy decisions. In line with this, $cor(y_t, \widehat{\eta}_t) = -0.08$ with a t-statistic of -1.30, implying that there is no statistically significant relationship, where y_t represents the CPI inflation in the Euro Area.

Regarding lead-lag exogeneity, any estimate $\hat{v}_{t+h}$ is, by construction, not correlated with $\hat{\eta}_t$. Specifically, any information contained in $\hat{\eta}_t$ and included in equation (4) is also present in $\widehat{\Delta r_t}$, where the latter is orthogonal to $\hat{v}_{t+h}$. Furthermore, the information set in x_t used to estimate β_h includes lags of y_t, ensuring lag-exogeneity.

3.2.2 Smooth Transition Autoregressive Model

Unlike the previous model specification based on Laumer and Schaffer (2025), the potential nonlinearity between global supply chain disruptions and the monetary policy transmission mechanism will be examined using a Smooth Transition Autoregression (STAR) model (Granger and Terasvirta, 1993):

$$\begin{aligned} y_{t+h} &= (1 - F(z_{t-1}))[\widehat{\Delta r_t}\beta_h^1 + x_t\gamma_h^1] + F(z_{t-1})[\widehat{\Delta r_t}\beta_h^2 + x_t\gamma_h^2] + v_{t+h} \\ F(z_t) &= \frac{e^{-\gamma z_t}}{1 + e^{-\gamma z_t}} \quad for \quad \gamma > 0 \end{aligned} \tag{5}$$

[23] I selected the number of lags as in the first model specification inspired by Laumer and Schaffer (2025). The most common suggested number of lag was 1, as also suggested in the Hernández et al. (2024) analysis. The results, however, are robust to the number of lags included.

The model framework presented above closely aligns with that of Ramey and Zubairy (2014), who employ a similar approach to analyze the government spending multiplier in different economic conditions. The key distinction lies in their choice of transition function $F(.)$, which in Ramey and Zubairy (2014) is defined as a simple indicator function, making it discrete. However, such a discrete function would not be suitable for my analysis, as it fails to align with the continuous nature of my state variable, the GSCPI.

For my purposes, I require a model capable of capturing both abrupt but temporary shifts and more gradual transitions in the GSCPI between high- and low-stress regimes. To achieve this, I adopt the above logistic transition function $F(.)$. The logistic function ensures that state changes occur smoothly, preventing minor fluctuations in z_t from triggering sudden discrete regime changes.

Each state is characterized by a distinct set of coefficients, as indicated by the superscripts in β_h^i, γ_h^i for $i = 1, 2$. Furthermore, the transition function $F(.)$ depends on the lagged cyclical component of the GSCPI state variable to avoid contemporaneous feedback between the state and the shocks as in Ramey and Zubairy (2014), Auerbach and Gorodnichenko (2012), Passos and de Melo Modenesi (2021) and Laumer and Schaffer (2025). The cyclical component is extracted using the Hodrick-Prescott filter, as applied in Auerbach and Gorodnichenko (2012) and Hernández et al. (2024). This filtering method ensures that the time series is recentered at zero, irrespective of the sample period. Following Hernández et al. (2024), in the baseline specification, I set $\lambda = 129600$ in the Hodrick-Prescott filter, as recommended by Ravn and Uhlig (2002), and $\gamma = 6$ in the logistic function. The parameter γ defines the intensity of the smoothing: higher values of γ mean that $F(z_t)$ stays longer near the limits $[0, 1]$, bringing the model closer to a discrete setting.[24]

As already said, I am interested in analyzing the response of accumulated consumer price inflation to an exogenous monetary policy intervention at time t. Following Hernández et al. (2024), I focus on a 24-month horizon, where β_h^i represents the accumulated CPI response at time $t + h$ under state $i = 1, 2$ of a 100 basis point increase in the monetary policy rate at time t, that is $\widehat{\Delta r_t} = 1$ percentage point.

3.3 Third Analysis

The third model specification is based on the work of Jordà et al. (2020), Hernández et al. (2024), and Di Giovanni et al. (2009). Overall, the estimated equations are not different from those in the second analysis. However, Denmark's unique monetary framework, characterized by free capital mobility and fixed exchange rate, provides an ideal environment for a faithful replication of Jordà et al. (2020)'s instrumental variable technique. In summary, there are additional and compelling reasons to support the validity of the previous instrumental variable empirical strategy. Therefore, I will utilize the Danish case to further corroborate and enhance the results.

As explained by Obstfeld et al. (2004), the concept that only two out of three policy objectives - fixed exchange rates, open capital markets, and autonomous monetary policy - can be simultaneously

[24] In the Appendix, I will verify that the results remain highly robust to the choice of parameter values.

achieved is known as the classic macroeconomic trilemma. The intuition is straightforward: when a country credibly and permanently pegs its exchange rate to that of a base country and allows free capital mobility, interest rate parity dictates that the domestic interest rate must align with the base countrys rate. This severely limits the monetary sovereignty of the pegged economy. In other words, when both the exchange rate is fixed and capital is freely mobile, monetary policy loses its effectiveness in pursuing domestic objectives. Any intervention in support of exchange parity then involves capital flows that exactly offset any monetary policy action threatening to alter domestic interest rates (Obstfeld and Taylor (1998)).

The works of Shambaugh (2004), Obstfeld et al. (2004), and Obstfeld et al. (2005) broadly support the trilemma and its implications. In pegged economies with open capital markets, short-term interest rates closely track the base countrys interest rate, indicating a significant loss of monetary autonomy compared to economies with alternative exchange rate regimes. Therefore, as Shambaugh (2004) notes, when a pegged countrys interest rates mirror those of the base country, its monetary policy is effectively decoupled from domestic economic conditions. In addition, there is no feedback loop from the local economy to the policy-setting process, since the base country's interest rates are determined without considering their impact on the pegged economy. This characteristic makes pegged economies a useful context for studying the effects of monetary policy. Jordà et al. (2020) leverages this intuition to empirically examine the monetary policy transmission mechanism in pegged economies. Specifically, they used exogenous fluctuations in pegged interest rates, driven by the unpredictable component of the base country's policy rate, to analyze the effects of monetary interventions using historical panel data dating back to 1870.[25]

Based on these considerations, and as done in the previous subsection, I use the unpredictable component of the federal funds rate (the base country rate):

$$\Delta r_t^* = \alpha_h^* + x_t^* \gamma_h^* + \eta_t \tag{6}$$

to isolate exogenous variations in the interest rate of Denmark:

$$\Delta r_t = \delta_0 + \widehat{\eta}_t \psi + x_t \delta + u_t \tag{7}$$

This exogenous variations will then be employed in a series of nonlinear local projections of the form:

$$\begin{aligned} y_{t+h} &= (1 - F(z_{t-1}))[\widehat{\Delta r_t}\beta_h^1 + x_t\gamma_h^1] + F(z_{t-1})[\widehat{\Delta r_t}\beta_h^2 + x_t\gamma_h^2] + v_{t+h} \\ F(z_t) &= \frac{e^{-\gamma z_t}}{1 + e^{-\gamma z_t}} \quad for \quad \gamma > 0 \end{aligned} \tag{8}$$

to investigate potential nonlinearities in the monetary policy transmission mechanism in relation to global supply chain disruptions.

[25] As discussed in the previous section, the European monetary policy rate closely tracked the US rate, making it a relevant instrument. Consequently, one could argue that the following considerations are not limited to pegged economies alone. Simply, in similar economies like Denmark, there are additional justifications for this empirical strategy. Therefore, the Danish case will be utilized to further corroborate and strengthen the results.

Denmarks monetary framework, characterized by free capital mobility and a fixed exchange rate, makes it an ideal case to apply this econometric methodology. Although the Danish krone is pegged to the euro, which would naturally serve as its base currency, I use the federal funds rate instead. This choice enhances the validity of the exclusion restriction while maintaining the relevance and exogeneity assumptions.

To be precise, as noted in Jordà et al. (2020), the validity of the instrument - and consequently the ability to draw causal inferences - depends on the following assumptions:

Assumption 1. *(Relevance and Exogeneity). I assume:*

- *Relevance:* $L(\Delta r|x, z; q = 1) \neq L(\Delta r|x; q = 1)$,
- *Exogeneity:* $L(y_j|x, \Delta r, z; q = 1) = L(y_j|x, \Delta r; q = 1)$ *for* $j = 0, 1$

where, for example, $L(\Delta r|x, z)$ *refers to the linear projection of* Δr *on* x *and* z *(the instrument).*

Note that I condition only on $q = 1$, indicating pegged economies; in other words, the above assumption needs to hold solely for the subpopulation of pegs, as implied by the trilemma-based application.

It is important to note that identification relies also on the exclusion restriction - the assumption that base country interest rates affect pegged economies solely through the interest rate channel. I treat the exclusion restriction as a subset of the broader exogeneity assumption. Specifically, if the exclusion restriction fails, then exogeneity also fails. However, a failure of exogeneity does not necessarily imply a violation of the exclusion restriction, as exogeneity can also be compromised by other factors.

Regarding the relevance condition, the instrument is statistically significant at the 1% level, exhibiting a positive coefficient of 0.12. As for the exogeneity condition, I argue that if it holds when using the euro area interest rate as the base rate, it should hold *a fortiori* when using the federal funds rate instead, given the fewer economic ties between the US and Denmark.[26]

From an economic perspective, a violation of the exclusion restriction could occur if base country interest rates influence domestic outcomes through channels other than domestic interest rate movements. These additional influences are commonly referred to as spillover effects, which may arise if base rates act as proxies for global factors affecting multiple economies simultaneously. However, for such spillovers to pose a significant concern, they would need to persist despite the inclusion of controls for global real GDP growth - intended to capture common business cycle fluctuations - and adjustments for both base and domestic economic conditions.

Moreover, I argue that using the federal funds rate as the base measure further reduces the likelihood

[26] As Shambaugh (2004) observes, when a pegged country's interest rates closely follow those of the base country, its monetary policy becomes effectively detached from domestic economic conditions. Furthermore, there is no feedback loop between the local economy and the policy-setting process, as the base country sets its interest rates without considering their effects on the pegged economy. This argument is even more relevant when the federal funds rate serves as the base rate.

of such spillover effects due to the relatively weaker economic linkages between the U.S. and Denmark. In support of this reasoning, Di Giovanni et al. (2009) employ a similar empirical strategy to estimate the effects of monetary policy on output growth. They exploit quasi-experimental variation in interest rates generated by the adherence of European countries to Germany's interest rates, with Germany serving as an 'anchor' country within a fixed exchange rate system that characterized many European economies during the period 1973-1998 following the collapse of the Bretton Woods system. Similar to my case, their estimation strategy is based on the assumption that the instrument exerts no direct effects on the domestic economy beyond the interest rate channel. However, given the strong trade and financial linkages among European economies, two related potential concerns arise with this instrumental variable approach. First, Germany's interest rate can directly affect the domestic economy of the follower countries beyond its transmission through domestic interest rates. This concern is particularly relevant for smaller countries more dependent on trade with Germany, which could experience adverse effects from a contraction in German demand following an interest rate hike. Secondly, this implies that output and inflation innovations are likely to be correlated across countries, potentially biasing the IV estimates due to cross-country economic interdependencies.

The use of federal fund rates makes this issue less likely in my analysis. To illustrate this point more clearly, consider the following simplified system of simultaneous linear equations:

$$\begin{aligned} y_t &= \alpha_0 + \theta i_t + u_t \\ i_t &= \beta_0 + \beta_1 z_t + \eta_t \\ y_t^* &= \alpha_0^* + \theta^* z_t + u_t^* \end{aligned} \tag{9}$$

where y_t is Danish inflation, i_t is the Danish interest rate, z_t is the base country interest rate, and variable with asterisks are base country variables. A failure of the exclusion condition occurs if:

$$u_t = \delta u_t^* + \omega_t \tag{10}$$

where I assume that ω_t and z_t are uncorrelated. However, z_t and u_t^* are inherently correlated, since the base country's interest rate is determined in response to its own economic conditions. If $\delta \neq 0$, the IV strategy becomes invalid due to the presence of correlated shocks and a direct effect mechanism. Intuitively, the stronger the economic ties between the base country and the pegged economy, the more likely it is that $\delta \neq 0$, indicating the existence of correlated shocks and potential direct effects beyond the interest rate channel. Di Giovanni et al. (2009) apply similar reasoning and find that the difference between the OLS and IV estimates is smaller for countries with a higher trade-to-GDP ratio with Germany. To the extent that IV cleans the estimates from its endogenous components, this suggests that stronger trade links with the base country may reduce the effectiveness of the IV strategy due to increased exposure to correlated economic shocks.

Using the federal funds rate as the base measure should help mitigate this problem, since Denmark's trade share with Europe in 2022 was 55.1%, compared to just 10.1% with the United States.

4 Results

In this section, I present my results, structured into two parts. The first part reports the findings of the econometric model based on Laumer and Schaffer (2025), while the second presents the results of the second and third analyses, focusing on the accumulated CPI response.

4.1 First Analysis - Results

Bringing together the findings from both the Euro Area and individual country-level analyses, I find neither conclusive evidence to fully corroborate nor sufficient grounds to reject Laumer and Schaffer (2025)'s amplification effect. The results remain mixed: industrial production exhibits a modest amplification effect, while CPI shows an even weaker one. No statistically significant differences emerge for retail trade, despite the Euro Area-level analysis suggesting an economically meaningful IRF with a significant amplification effect - an outcome not confirmed by the more granular country-level analysis. Finally, the response of the unemployment rate appears to be largely unaffected by the state of global supply chains. Extending the analysis to include COVID-19 data does not materially alter the key conclusions. The only notable divergence is the emergence of a modest amplification mechanism for unemployment, although it generally remains statistically insignificant. This contrasts with Laumer and Schaffer (2025), who report a more pronounced amplification effect when incorporating post-COVID-19 observations.

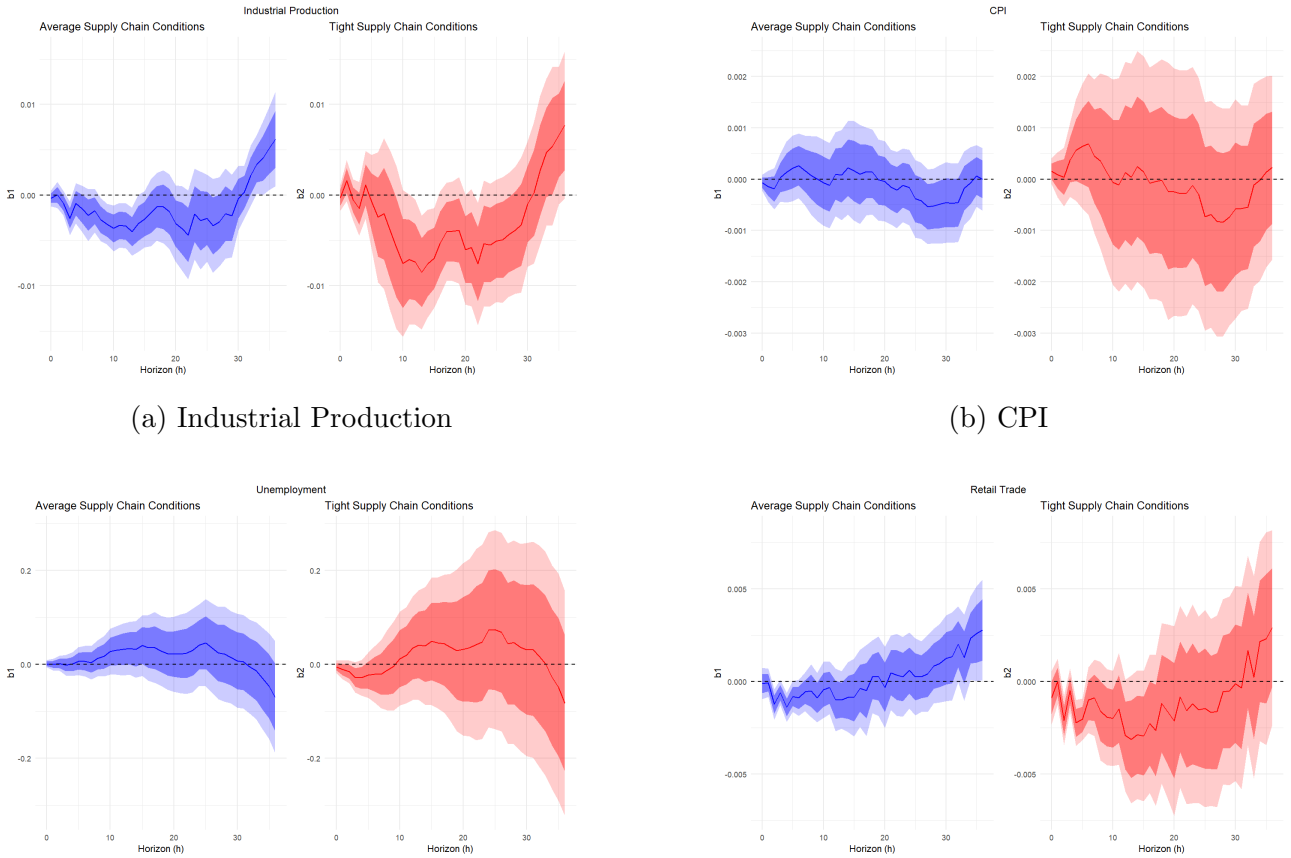
4.1.1 Euro Area

The Euro Area state dependent IRFs for industrial production, CPI inflation, unemployment rate, and retail trade, together with their 68% and 90% confidence intervals, are presented in Figure 5. They are based on an adapted version of the econometric methodology of Laumer and Schaffer (2025), using data from January 2001 to December 2019.

The IRFs generally align with predictions from standard economic theory. In general, a strong amplification mechanism, as identified in Laumer and Schaffer (2025), does not emerge clearly, with mixed evidence remaining. Although a weak amplification effect is observed for industrial production and a stronger one for retail trade, no statistically significant differences are found between the two states for the unemployment rate and the CPI. These latter variables show no significant response under either condition, suggesting that they remain largely neutral to monetary policy actions. Also, a moderate price puzzle emerges when focusing on the first horizons.

The amplification effect is observable in Figure 6, which illustrates the key coefficients central to my analysis, $\{\delta_0^h\}_{h=0}^H$. These coefficients capture potential state dependencies by quantifying the additional impact of a monetary policy shock when global supply chain pressures are elevated - specifically, one standard deviation above their average level.

Only the coefficients for industrial production, and especially retail trade, indicate a modest yet statistically significant amplification mechanism in line with Laumer and Schaffer (2025). In contrast, no clear pattern emerges for the unemployment rate and CPI.



(a) Industrial Production

(b) CPI

(c) Unemployment Rate

(d) Retail Trade

Figure 5: Euro Area state-dependent impulse responses. Macro variables (2019)

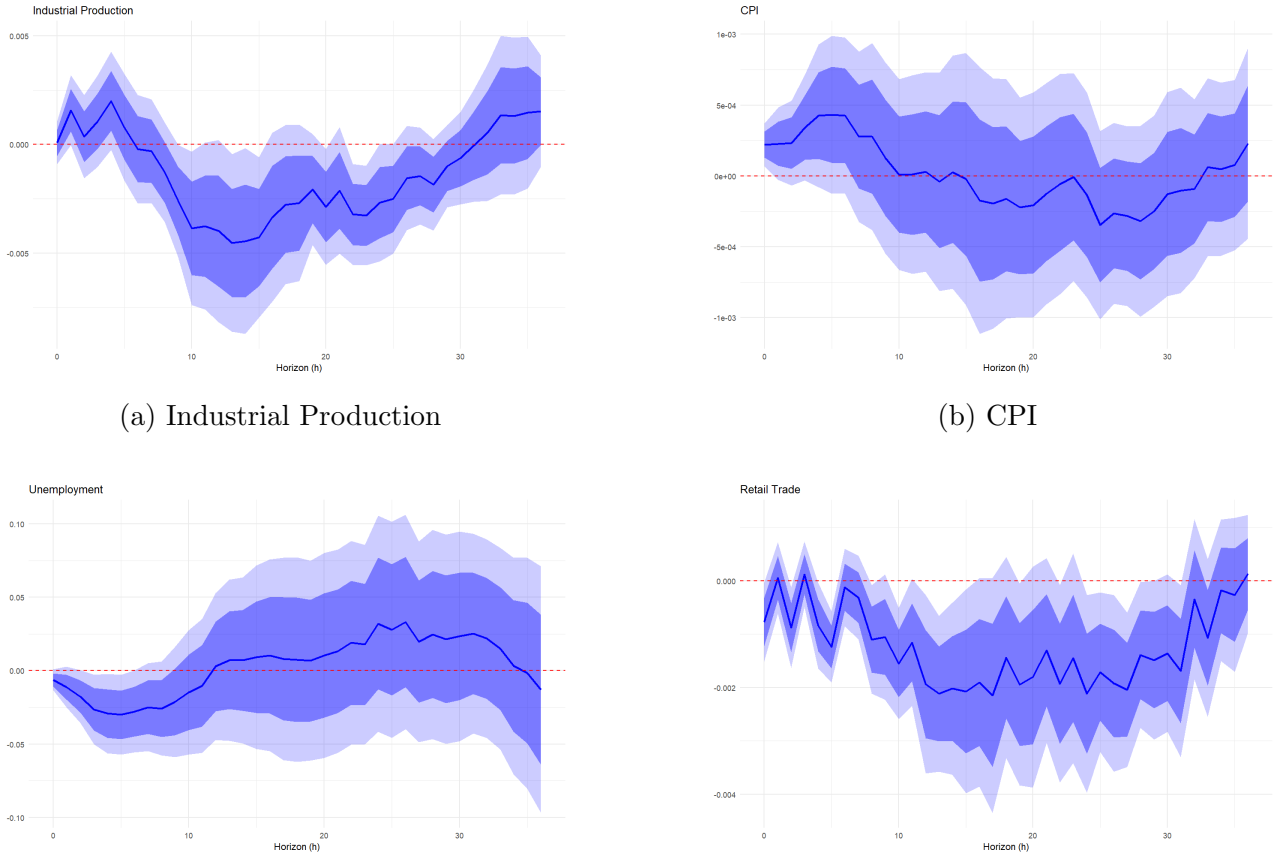
4.1.2 Euro Countries

I applied the same econometric methodology to data from the largest Euro Area economies - Germany, France, Italy, and Spain - to explore cross-country heterogeneity, test the robustness of the results, and gain additional insights. Similarly to the Euro Area analysis, no clear and strong amplification mechanism is revealed, as identified in Laumer and Schaffer (2025).

Inspecting Figure 7, industrial production exhibits economically meaningful impulse response functions, with a stronger amplification effect in Germany and Italy, but weaker in France and Spain.

With the exception of Spain and Germany under tight supply chain conditions, the individual countries' CPI impulse response functions appear to resolve the price puzzle observed at the Euro Area level in the first horizons. Overall, the IRFs display the expected sign, with France exhibiting the most sensitive inflation rate to monetary shocks and with some evidence of monetary neutrality, particularly in Germany under average supply chain conditions. Additionally, there appears to be a weak amplification mechanism for CPI.

The IRFs for unemployment are generally economically meaningful, except in Italy and, particularly, Spain, where unemployment appears to be money neutral. In Spain, the unemployment IRF exhibits an economically insignificant downward trend over longer horizons. This suggests that unemployment in the region is likely driven by specific economic factors that my econometric setup cannot fully cap-



(a) Industrial Production

(b) CPI

(c) Unemployment Rate

(d) Retail Trade

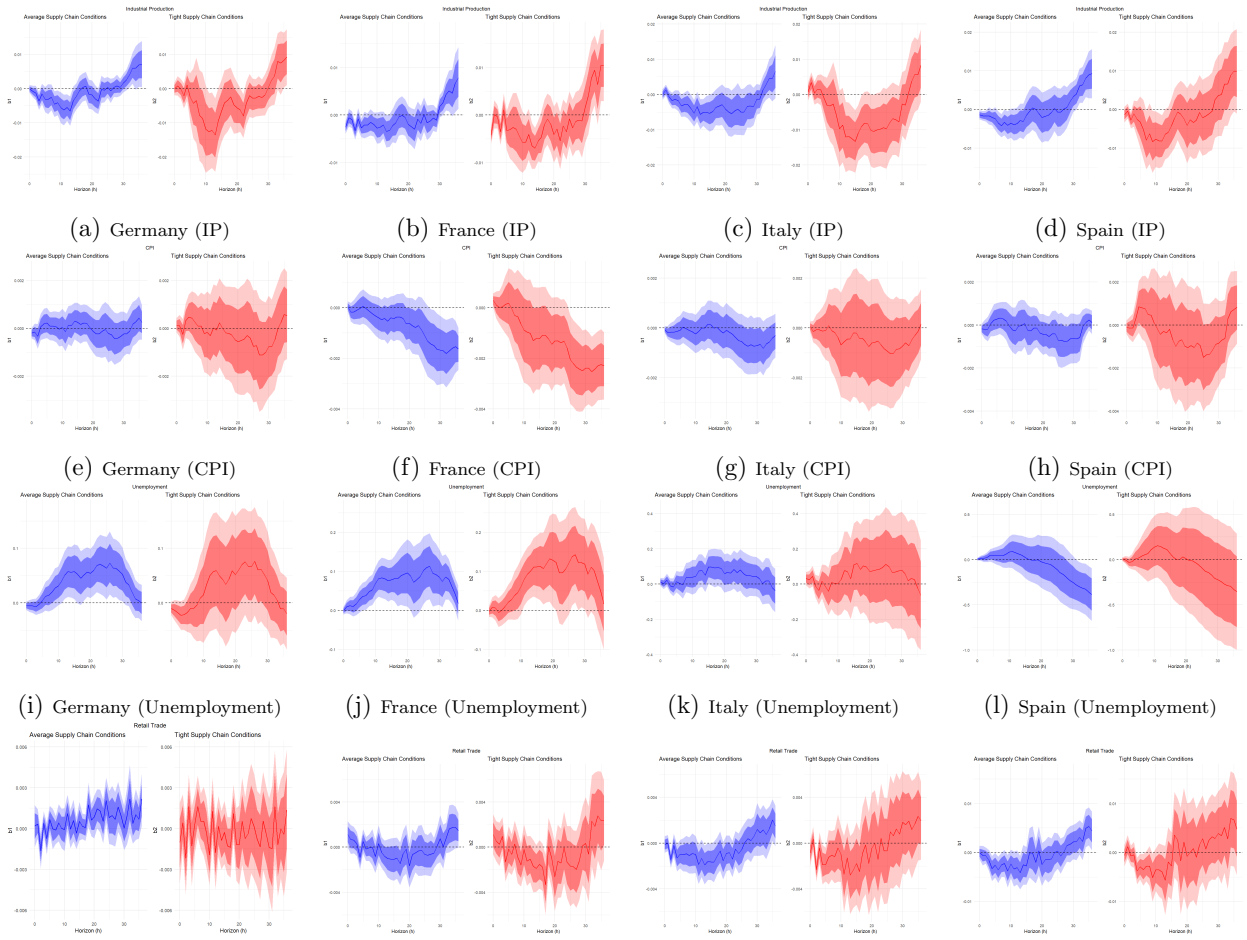
Figure 6: Euro Area differential responses. Macro variables (2019)

ture. In France and Germany, the unemployment response aligns with expectations. Overall, there is no evidence of an amplification effect for unemployment.

Finally, the IRFs for retail trade exhibit significant volatility, particularly in Germany, while being somewhat more stable in other countries, where they generally display the expected negative sign at shorter horizons. In general, the responses are not statistically significant and no clear amplification effect is observed.

To further investigate state-dependent differences, Figure 8 presents the coefficients $\{\delta_0^h\}_{h=0}^H$, which capture potential state dependencies. This figure reinforces the findings discussed so far. A modest amplification mechanism is observed for industrial production, with an even weaker effect for CPI, where across most horizons the difference between the two states is not statistically significant at the 68% level, with the exception of France. No clear amplification effect emerges for the unemployment rate. Germany exhibits a pattern somewhat opposite to amplification, Italy shows virtually no difference between the two states, and Spain and France display a very weak amplification effect, which is generally not statistically significant at the 68% level.

Finally, the response of retail trade remains highly erratic, with no statistically significant differences between the two states. The only exception is Germany, where a weak amplification mechanism appears at longer horizons.



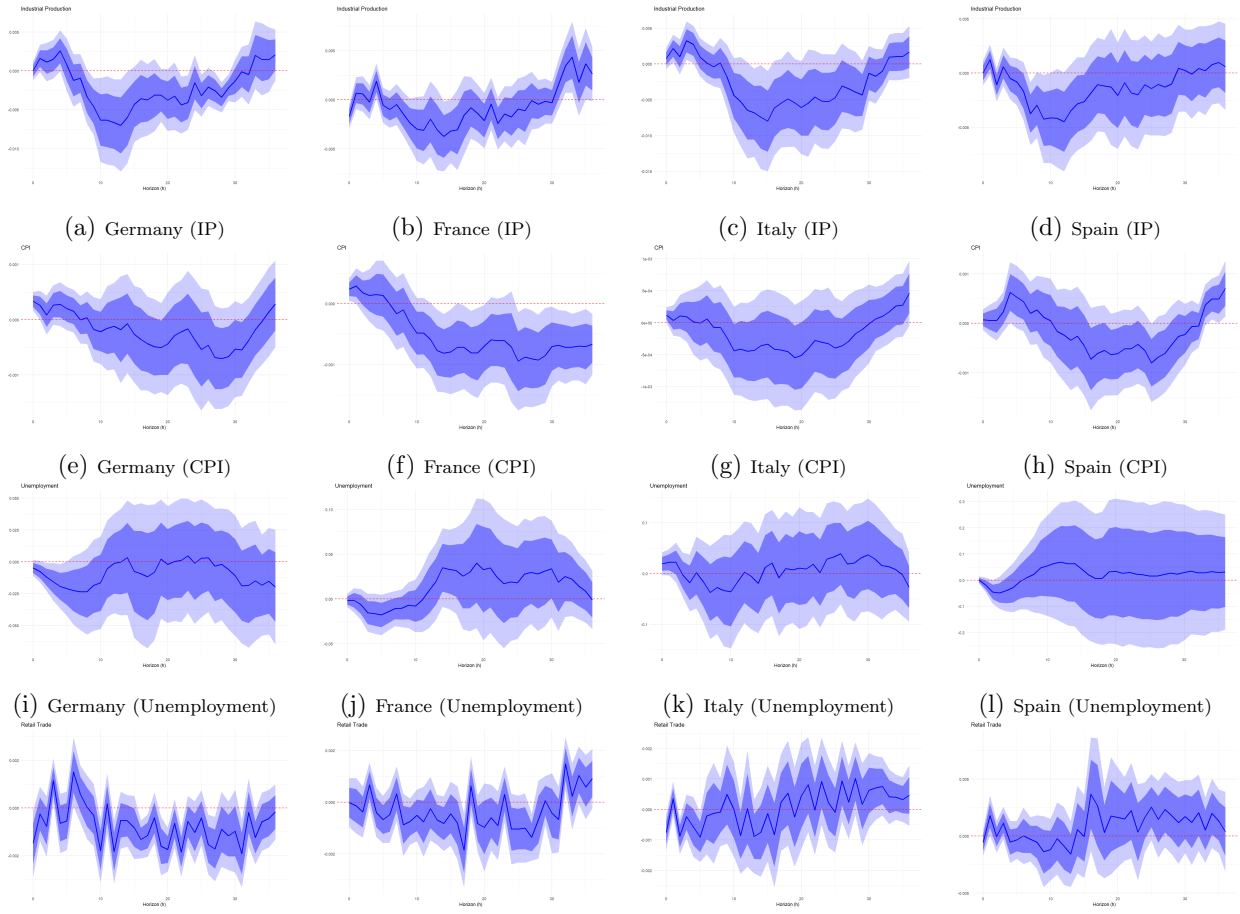
(a) Germany (IP) (b) France (IP) (c) Italy (IP) (d) Spain (IP)

(e) Germany (CPI) (f) France (CPI) (g) Italy (CPI) (h) Spain (CPI)

(i) Germany (Unemployment) (j) France (Unemployment) (k) Italy (Unemployment) (l) Spain (Unemployment)

(m) Germany (Retail Trade) (n) France (Retail Trade) (o) Italy (Retail Trade) (p) Spain (Retail Trade)

Figure 7: Euro countries state-dependent impulse responses. Macro variables (2019)



(a) Germany (IP) (b) France (IP) (c) Italy (IP) (d) Spain (IP)

(e) Germany (CPI) (f) France (CPI) (g) Italy (CPI) (h) Spain (CPI)

(i) Germany (Unemployment) (j) France (Unemployment) (k) Italy (Unemployment) (l) Spain (Unemployment)

(m) Germany (Retail Trade) (n) France (Retail Trade) (o) Italy (Retail Trade) (p) Spain (Retail Trade)

Figure 8: Euro countries differential responses. Macro variables (2019)

4.1.3 Covid Extension

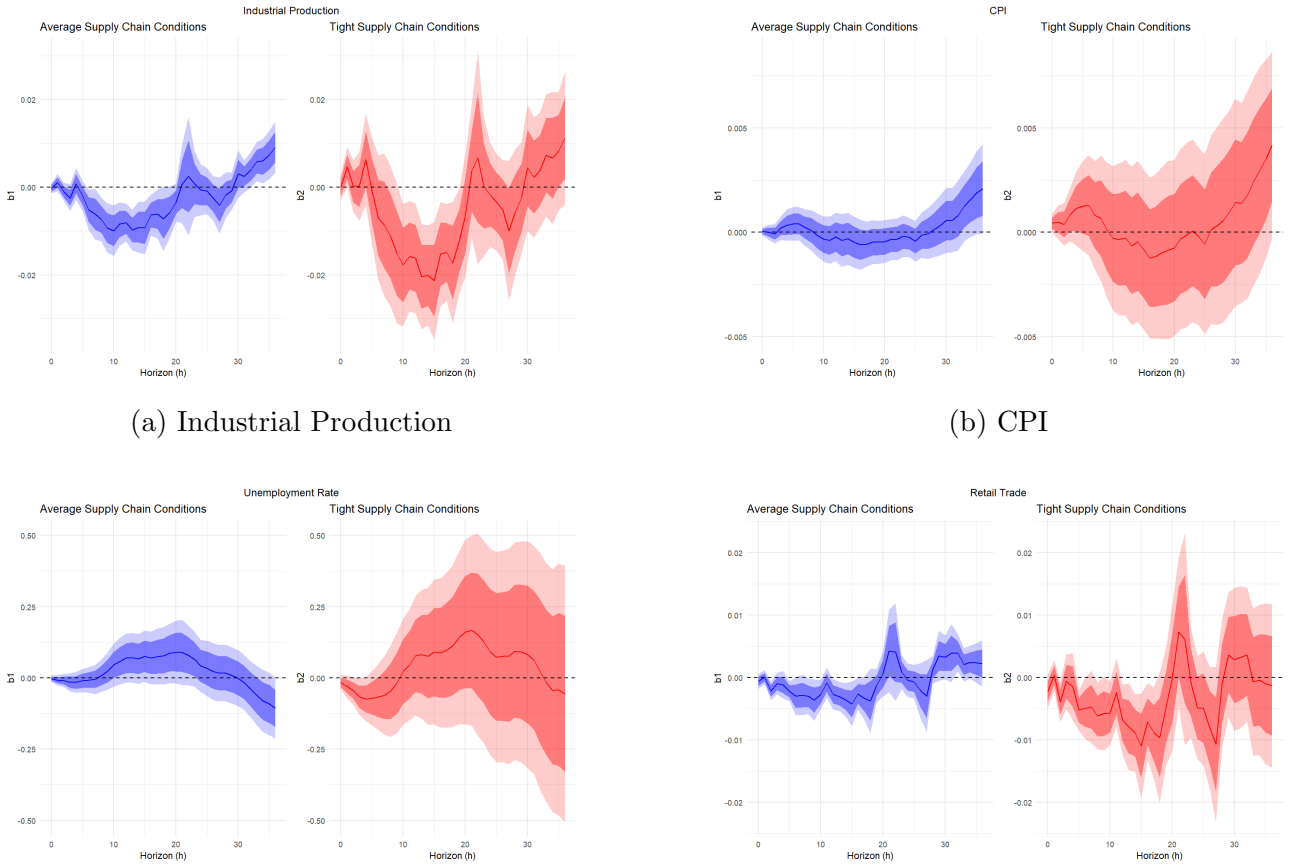
As highlighted by Ng (2021), many economic time series may have experienced structural breaks in their data-generating processes during the pandemic. The COVID-19 period also introduced numerous confounding factors, such as government restrictions on mobility and business operations, which complicate empirical analysis. Consequently, and in line with the baseline model specification of Laumer and Schaffer (2025), I restrict my main analysis to a pre-pandemic sample covering January 2001 to December 2019. However, given the profound impact of supply chain disruptions on the post-pandemic economy, it is important to test the robustness of the results by incorporating more recent observations. To this end, this section extends the baseline analysis to include data up to December 2022, the last full year for which all variables have available observations.

To address the substantial volatility, outliers, and potential structural breaks triggered by the COVID-19 pandemic, the literature offers several approaches. Ng (2021) suggests incorporating additional COVID-related controls - such as hospitalizations, positive cases, and deaths - to "decovid" the data. Following this approach, Laumer and Schaffer (2025) include national COVID-19 cases and deaths as controls for potentially altered dynamics in the post-pandemic period, finding that their inclusion has minimal impact on the estimation results. Meanwhile, Lenza and Primiceri (2020) propose explicitly modeling the change in shock volatility to account for the exceptionally large macroeconomic innovations observed during the pandemic.

In this study, I take a different approach by including monthly dummies for each post-COVID-19 month in my sample, from March 2020 to December 2022. Estimating the baseline model without adjustments using observations through December 2022 yields highly volatile or economically insignificant IRFs, particularly for CPI. However, the inclusion of monthly dummies helps filter out pandemic-induced distortions, producing IRFs that align more closely with the baseline analysis.

Figure 9 presents the Euro Area state-dependent IRFs. Overall, the results align with the pre-pandemic analysis conducted at the Euro Area level, with retail trade and industrial production displaying a modest amplification effect. A notable difference compared to the pre-pandemic findings is a slightly heightened sensitivity of the unemployment rate to monetary policy when global supply chains are under stress. However, as in the pre-pandemic analysis, no significant differences are observed between the two states in the response of the CPI, still displaying a moderate price puzzle for the first horizons. Note also that the CPI response exhibits an economically meaningless upward trend at longer horizons. This puzzling result suggests that the dummy approach may not fully account for all COVID-19-related distortions, potentially leading to the observed anomaly. The subsequent analysis at the individual country level will help determine whether this trend is a pervasive feature of the broader analysis or if it is specific to the Euro Area sample.

To further investigate state-dependent differences, Figure 10 presents the coefficients $\{\delta_0^h\}_{h=0}^H$. A visual inspection of the figure confirms the pre-pandemic findings. With the exception of a weak amplification effect for the unemployment rate - statistically insignificant at the 68% level and only becoming apparent after horizon 12 - the results remain consistent. Specifically, a weak amplification mechanism persists for industrial production and retail trade, while no statistically significant differ-



(a) Industrial Production

(b) CPI

(c) Unemployment Rate

(d) Retail Trade

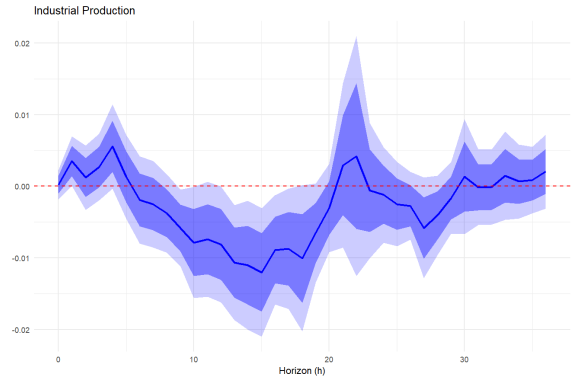
Figure 9: Euro Area state-dependent impulse responses. Macro variables (2022)

ences are observed for CPI.

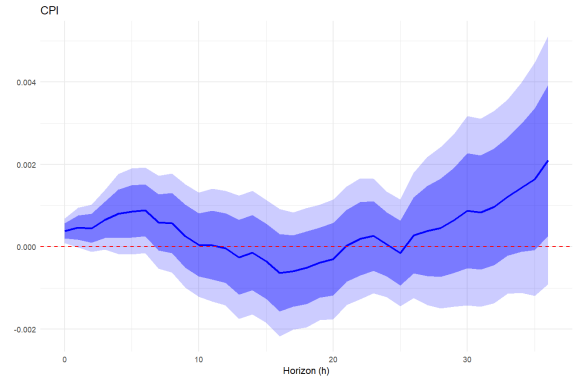
Figure 11 presents the results of the individual country analyses. Overall, the results align with the pre-pandemic evidence. Industrial production exhibits economically meaningful impulse response functions, consistent with the amplification mechanism identified by Laumer and Schaffer (2025). The IRFs for CPI all show the puzzling upward trend at longer horizons, even in France, despite its response remaining negative across all horizons. This supports the previous concern that the dummy approach may not fully control for all COVID-19-related distortions, potentially explaining this anomaly. Ignoring this puzzling behavior, the pre-pandemic findings hold: France exhibits the highest sensitivity of inflation to monetary shocks, while there is some evidence of monetary neutrality, particularly in Germany under average supply chain conditions. Additionally, a weak amplification mechanism for CPI appears across the samples considered.

The unemployment results are also broadly consistent with the pre-pandemic findings, with IRFs generally displaying economically meaningful patterns. However, exceptions emerge in Italy, where there is evidence of monetary neutrality and no clear state-dependent effects. A notable new finding is the presence of a weak amplification mechanism across all countries - except for Italy.

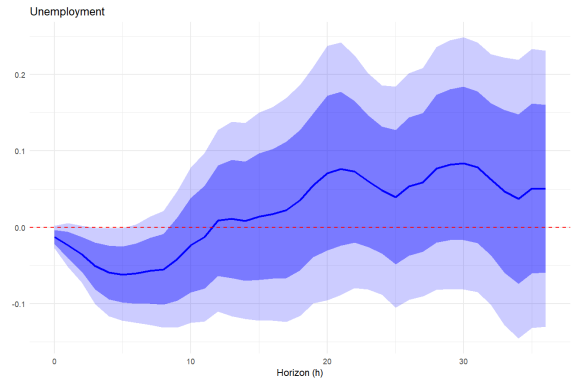
Finally, in line with the pre-pandemic analysis, the IRFs for retail trade remain highly erratic, difficult to interpret, and show no clear evidence of an amplification mechanism. To further investigate



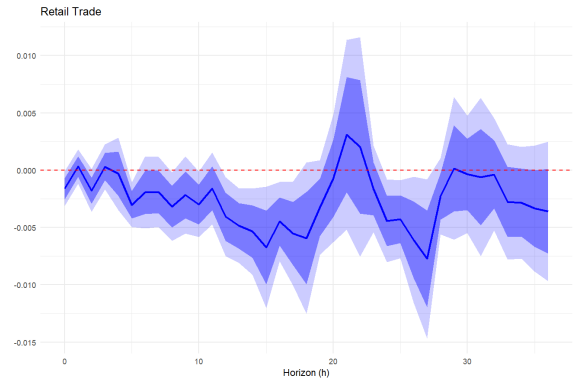
(a) Industrial Production



(b) CPI



(c) Unemployment Rate



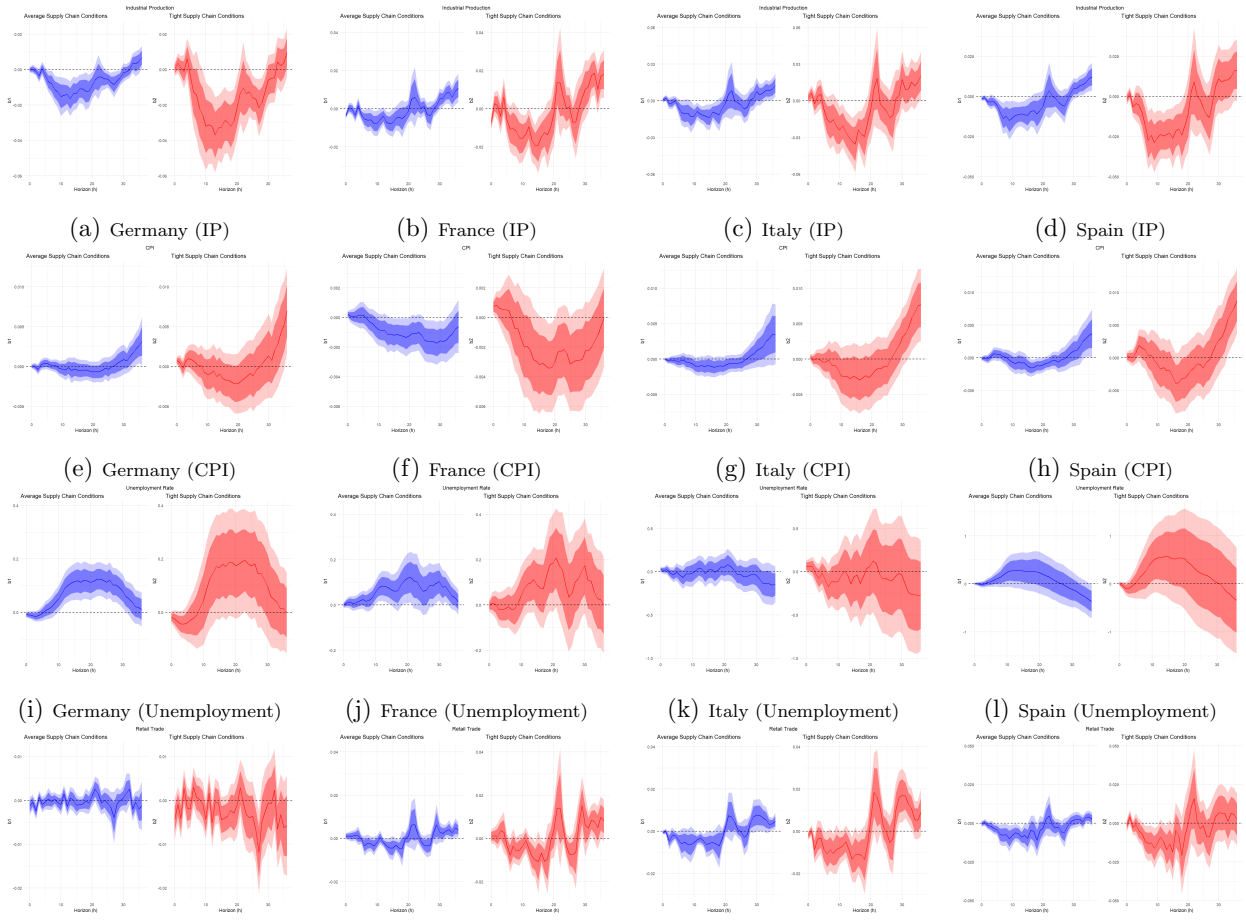
(d) Retail Trade

Figure 10: Euro Area differential responses. Macro variables (2022)

state-dependent differences, Figure 12 presents the key coefficients $\{\delta_0^h\}_{h=0}^H$. This figure reinforces the findings discussed so far. A modest amplification mechanism is observed for industrial production, with an even weaker effect for CPI, excluding the puzzling trends at longer horizons. Despite this anomaly, across most horizons - and with the exception of France, where the difference is more statistically significant - the disparity between the two states is only rarely statistically significant at the 68% level.

A deviation from the pre-pandemic results emerges in the case of unemployment. A very weak amplification effect is evident, though its strength varies across countries. In Germany, this effect is more pronounced at mid-range horizons and reaches statistical significance at the 68% level. In Spain, the amplification effect is more homogeneous across horizons, although it does not reach statistical significance. In France, it is nearly negligible, while in Italy, the results suggest somehow the opposite of an amplification effect, though this difference is also not statistically significant.

Finally, the response of retail trade remains highly erratic. A modest amplification effect is observed in Germany, whereas the other countries show mixed evidence. For the first half of the horizons, the amplification effect is weaker, whereas the second half follows a somewhat different pattern. However, in all cases, these effects generally lack statistical significance.



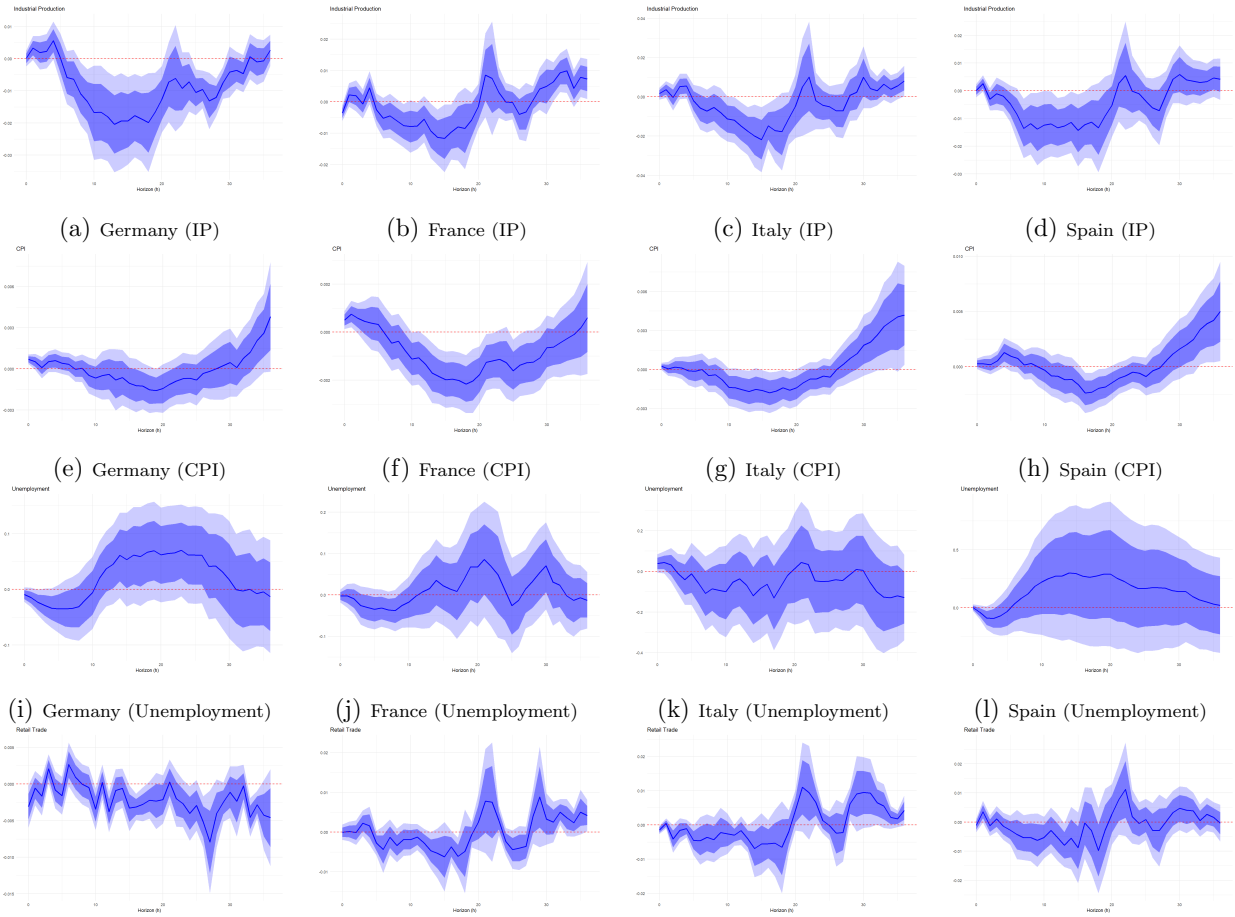
(a) Germany (IP) (b) France (IP) (c) Italy (IP) (d) Spain (IP)

(e) Germany (CPI) (f) France (CPI) (g) Italy (CPI) (h) Spain (CPI)

(i) Germany (Unemployment) (j) France (Unemployment) (k) Italy (Unemployment) (l) Spain (Unemployment)

(m) Germany (Retail Trade) (n) France (Retail Trade) (o) Italy (Retail Trade) (p) Spain (Retail Trade)

Figure 11: Euro countries state-dependent impulse responses. Macro variables (2022)



(a) Germany (IP) (b) France (IP) (c) Italy (IP) (d) Spain (IP)

(e) Germany (CPI) (f) France (CPI) (g) Italy (CPI) (h) Spain (CPI)

(i) Germany (Unemployment) (j) France (Unemployment) (k) Italy (Unemployment) (l) Spain (Unemployment)

(m) Germany (Retail Trade) (n) France (Retail Trade) (o) Italy (Retail Trade) (p) Spain (Retail Trade)

Figure 12: Euro countries differential responses. Macro variables (2022)

4.1.4 Connecting the Dots

Bringing together the results from both the Euro Area and individual country-level data, including and excluding COVID-19 observations, we find neither compelling evidence to fully confirm nor sufficient grounds to reject Laumer and Schaffer (2025)'s amplification effect. The evidence remains mixed: industrial production exhibits a modest amplification effect, while CPI displays an even weaker one. No statistically significant differences emerge for retail trade, despite the analysis at the Euro Area level that highlights an economically meaningful IRF with a significant amplification effect. However, this finding is not supported by the more granular analyses at the country level. Finally, the unemployment rate response appears largely unaffected by the state of global supply chains, with a weak but generally statistically insignificant amplification effect emerging in the COVID-19 extension.

Inflation results contrast with the findings of Hernández et al. (2024) and Andriantomanga et al. (2023), which provide evidence of a reduced ability of central banks to stabilize inflation amid global supply chain disruptions due to the relevance of second-round effects. Instead, my results are somehow closer to the conclusions of Laumer and Schaffer (2025) and Bai et al. (2024), suggesting an increased ability of monetary authorities to stabilize inflation during periods of global supply chain disruptions. This is likely due to the more credible monetary frameworks in the US and the Euro Area, which prevent inflation expectations from becoming unanchored, thereby limiting the second-round effects. This finding further underscores the benefits of maintaining a strong and credible monetary framework. However, the magnitude of the amplification mechanism is generally not statistically significant, and incorporating COVID-19 observations leads to puzzling results at longer horizons, likely due to unaccounted-for pandemic-related disturbances. This further underscores the need to analyze CPI responses under global supply chain disruptions using an alternative approach. The next section precisely does so.

On the other hand, the ambiguous results for real variables prevent a definitive conclusion on the economic costs of stabilizing inflation when global supply chains are under stress. While the modest amplification mechanism observed for industrial production aligns with Laumer and Schaffer (2025)'s findings, the absence of statistically significant differences for the unemployment rate and retail trade does not. Finally, none of the results supports Bai et al. (2024), who reports a reduced sensitivity of output to monetary policy shocks under global supply chain disruptions - echoing the seminal insights of Keynes (1940). These findings have important implications, as they help quantify the economic and social costs of stabilizing inflation in these exceptional circumstances. By shedding light on how global supply chain disruptions influence central banks' trade-offs, they contribute to a deeper understanding of the challenges monetary policymakers face during periods of heightened supply constraints. However, on this matter, my analysis does not yield a definitive answer.

4.2 Second and Third Analysis - Results

The previous section highlighted a modest amplification mechanism for industrial Production and a weaker one for CPI. Given the European Central Bank's primary mandate of maintaining price stability, I resorted to additional econometric models to examine more closely how inflation responds to

monetary policy interventions under different global supply chain conditions. Specifically, the focus is on the accumulated response of CPI inflation.[27] I analyze both the Euro Areas CPI inflation response and those of individual countries - Germany, France, Italy, and Spain - to assess the robustness of the results and explore potential heterogeneous effects. Finally, I leverage Denmark's unique monetary framework to further validate my results.

To assess the statistical significance of the state-dependent effect, I replicate the statistical tests conducted by Hernández et al. (2024), namely the Welch (1951) t-test for differences in two independent means. Following their approach, I report the test statistic for the null hypothesis that $\beta_h^1 = \beta_h^2$ for each h and refer to this as the point-by-point Welch test. I also present the p-value for the test of the null hypothesis $\sum_{h=1}^{H} H^{-1}\beta_h^1 = \sum_{h=1}^{H} H^{-1}\beta_h^2$ and refer to this as the mean Welch test.[28] Overall, my findings support the presence of weak to modest amplification mechanisms for CPI, with minimal heterogeneity across the analyzed samples. A modest amplification effect is observed in the Euro Area, France, Spain, and Italy, a weaker one in Denmark, and no state-dependent difference in Germany.

4.2.1 Euro Area

The accumulated impulse response of the Euro Area Consumer Price Index (log difference) to a monetary policy shock at time 0, along with its 90% and 68% confidence interval, is shown in the figure below. A modest amplification mechanism seems to emerge, as the accumulated impulse response of CPI under tight supply chain conditions is consistently lower than that under average supply chain conditions across all horizons.

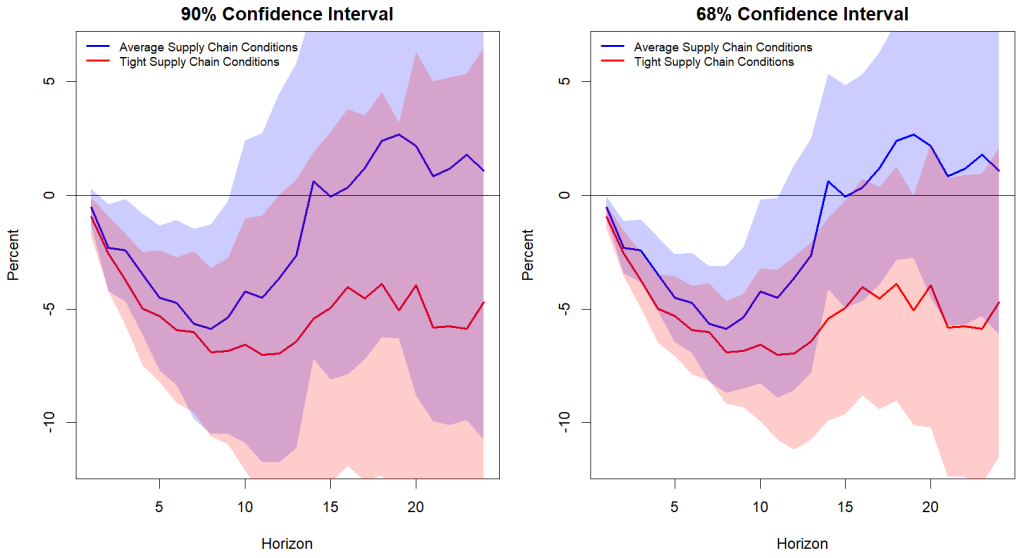


Figure 13: Accumulated Response of Consumer Price Index (log-difference)

[27] $100(ln(CPI_{t+h}) - ln(CPI_t))$

[28] These statistical tests are only provided to give a preliminary and intuitive understanding of the statistical significance of the results. A more robust statistical test should carefully consider the distribution of the residuals which most likely are strongly correlated with one another.

To quantify the statistical significance of the amplification effect, Figure 14 presents the point-by-point Welch t-test results, accompanied by a 90% confidence interval. The dotted lines indicate the bounds at ±1.6:

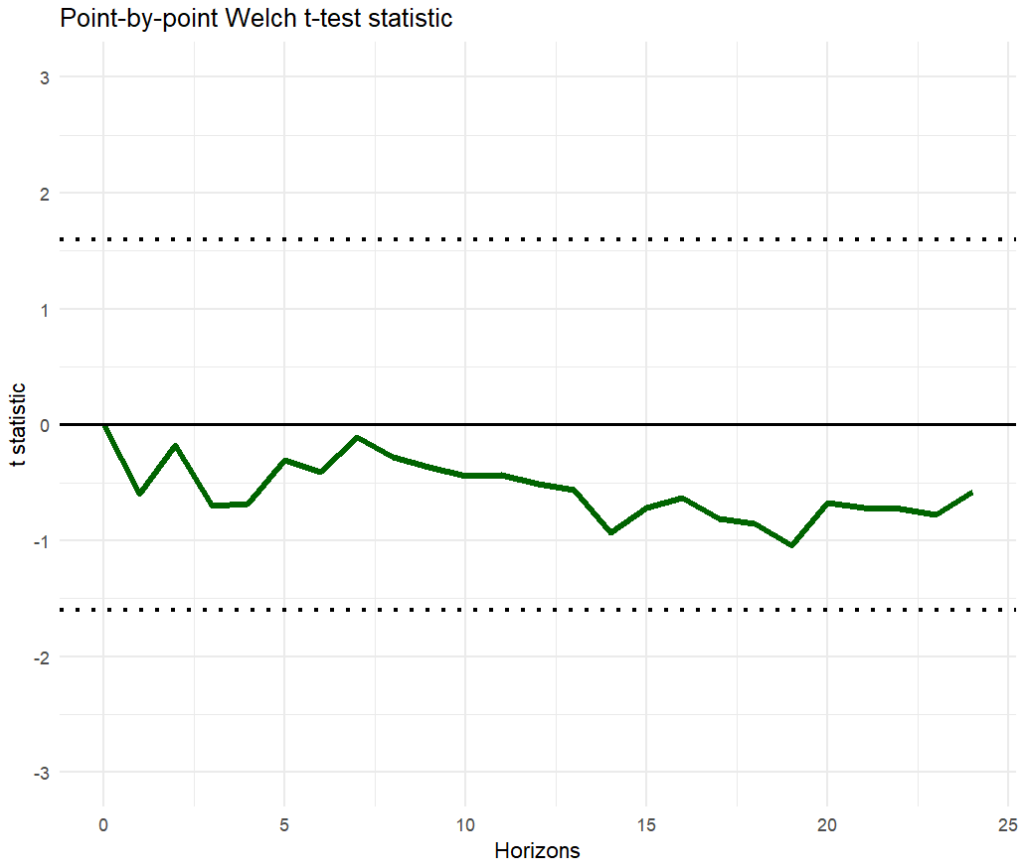
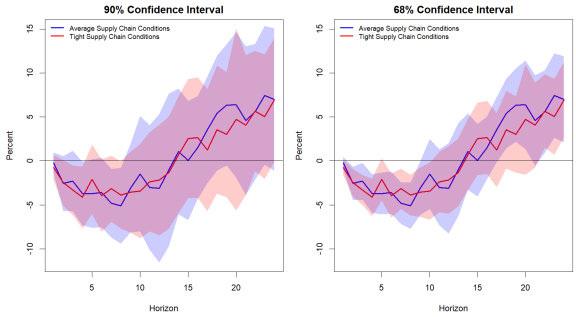


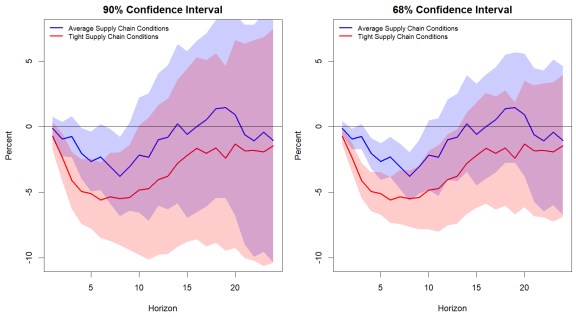
Figure 14: Point-by-point Welch t-test statistic (Euro Area)

Despite the point-by-point differences never reaching statistical significance at the 90% level, the mean Welch test yields a p-value close to zero thus suggesting that the IRFs are on average different. This apparent contradiction arises from the unexpectedly high standard errors of the Euro Area state-dependent IRFs reported earlier. However, in general, the evidence suggests a modest amplification mechanism. I now examine whether this finding holds across individual countries.

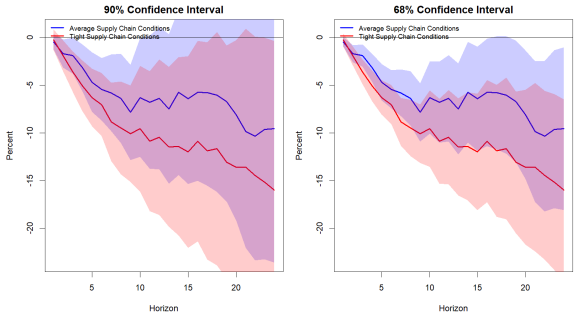
4.2.2 Euro Countries



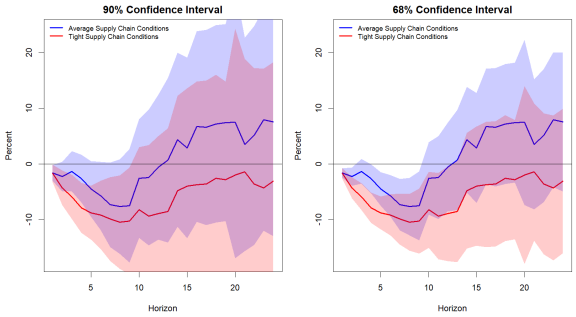
(a) Germany



(b) France



(c) Italy

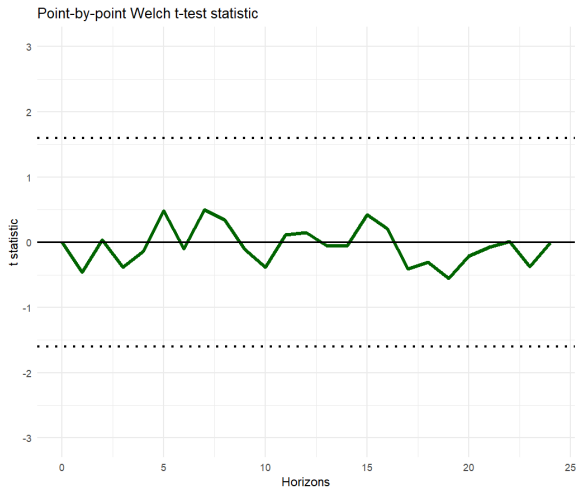


(d) Spain

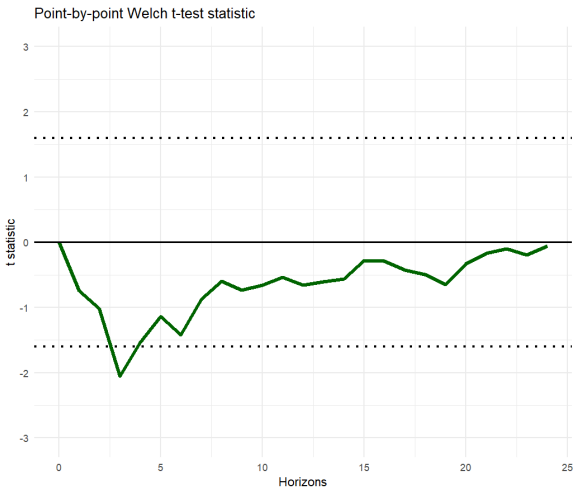
Figure 15: Accumulated Response of Consumer Price Index (log-difference)

The results of the analysis for individual European countriesGermany, France, Italy, and Spainare presented in Figure 15. Overall, the accumulated IRFs, along with the 90% and 68% confidence intervals, confirm the presence of a weak/modest amplification mechanism.

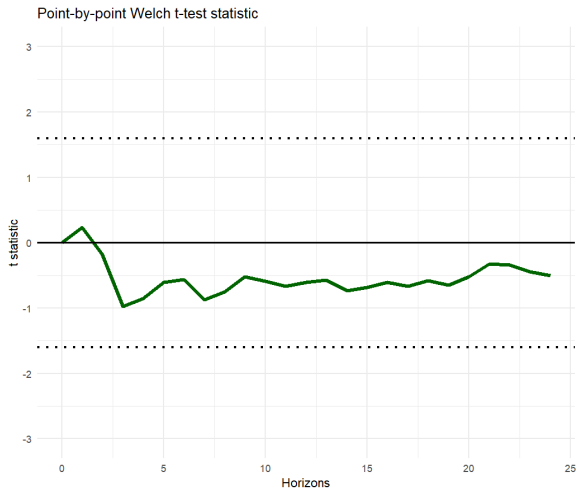
This is more evident when analyzing the results of the statistical test quantifying the significance of the amplification mechanism (Figure 16): with the exception of Germany, where no amplification mechanism is observed and the mean Welch test provides a p-value of 0.78, a modest amplification effect is present in the other regions. The point-by-point Welch t-tests are rarely statistically significant at the 90% level, with the sole exception of horizon 3 in France. However, similar to the Euro Area case, the mean Welch test yields very small p-values, close to zero, providing evidence in favor of an amplification mechanism. The discrepancy between the point-by-point results and the mean test arises from the high state-dependent standard errors, which lead to low t-statistics. Overall, the evidence supports the presence of a weak amplification mechanism.



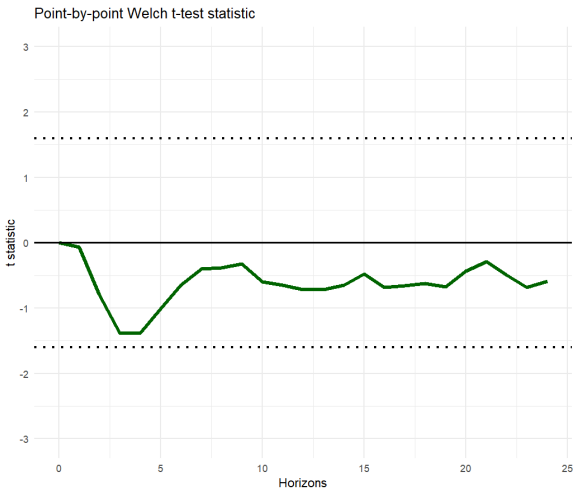
(a) Germany



(b) France



(c) Italy



(d) Spain

Figure 16: Point-by-point Welch t-test statistic (Individual Countries)

4.2.3 Denmark Case

This section presents the results of the third analysis, which further investigates the response of CPI under different supply chain conditions, with a particular focus on Denmark. The results are illustrated in Figure 17, which shows the accumulated IRFs along with their 90% and 68% confidence intervals.

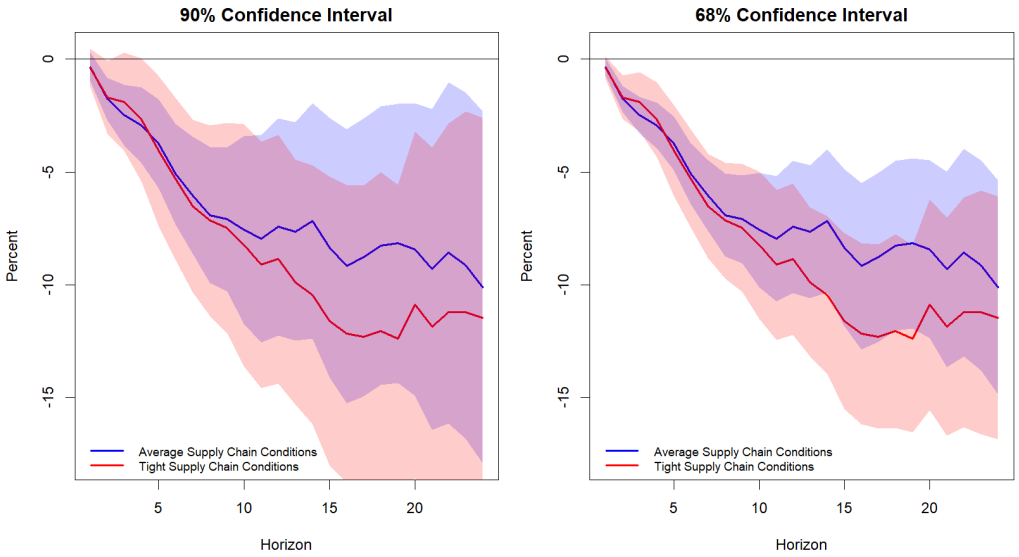


Figure 17: Accumulated Response of Denmark's Consumer Price Index (log-difference)

The amplification mechanism in Denmark appears weaker. As expected, the point-by-point Welch t-test (Figure 18) never reaches statistical significance due to the exceptionally high state-dependent standard errors. However, unlike in previous cases, the mean Welch t-test yields a significantly higher p-value of 0.10, suggesting a markedly weaker amplification mechanism.

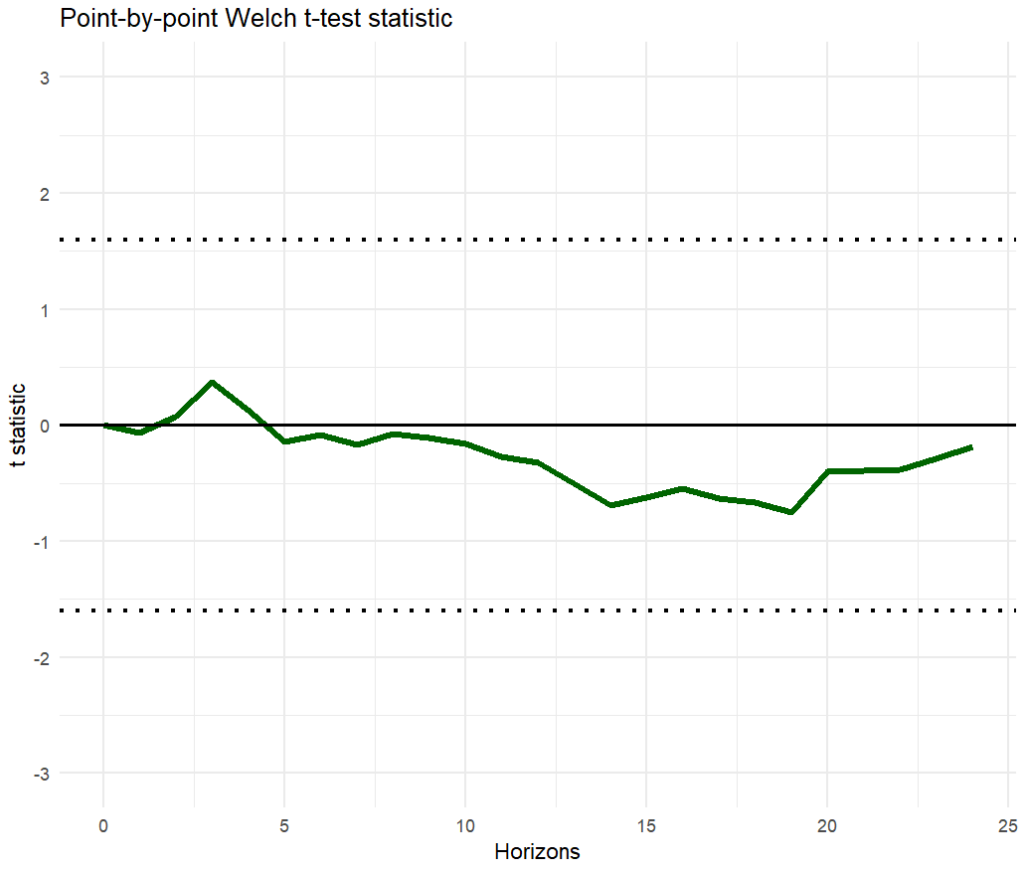


Figure 18: Point-by-point Welch t-test statistic (Denmark)

4.2.4 Connecting the Dots

The weak amplification effect in the CPI response observed in the first analysis, combined with the ECB's primary mandate of maintaining price stability, motivated further investigation. The last two sections examined the accumulated IRF of CPI under varying supply chain conditions and identified weak to modest amplification effects, qualitatively in line with the findings of Laumer and Schaffer (2025). This suggests an enhanced ability of central banks to maintain price stability when global supply chains are under stress. The next section explores the underlying economic mechanisms driving these results and seeks to provide a theoretical explanation.

5 Behind the Amplification Mechanism

My analysis does not identify a clear broad amplification mechanism, in contrast to the findings of Laumer and Schaffer (2025). Instead, the initial results suggest a modest amplification effect for industrial production and a weaker one for CPI, with no statistically significant differences for the unemployment rate and retail trade. The second and third analyses, which employ an instrumental variable approach, reinforce the CPI findings from the first analysis, highlighting the presence of weak to modest amplification effects. In this section, I try to interpret these results by testing the hypothesis proposed by Laumer and Schaffer (2025) using data from the Euro Area.

5.1 The Theoretical Mechanism

Laumer and Schaffer (2025) suggest that the stronger amplification effect observed in their analysis is driven by an intensified role of the credit channel. To illustrate this mechanism, consider a simplified representation of the credit channel framework from Bean et al. (2002). Due to credit market imperfections, the interest rate faced by firms, R_t, is determined by the risk-free rate, R_t^*, and an external finance premium, $f_t\left(\frac{D_t}{E_t}\right)$, which depends on the financial structure of the firm:

$$R_t = R_t^* + f_t\left(\frac{D_t}{E_t}\right) \qquad (f' > 0) \tag{11}$$

Here, R_t^* represents the central banks policy rate, while the external finance premium is an increasing function of the firms debt-to-equity ratio, reflecting its financial health. The key idea is that changes in the risk-free rate affect not only borrowing costs directly but also indirectly influence the firms financial position. In other words, an increase in R_t^* increases the cost of borrowing by more than a one-to-one ratio because it triggers an increase in the external finance premium through a rise in the $\frac{D_t}{E_t}$ ratio. This effect occurs because many firms finance their inventories and working capital with short-term debt, meaning higher interest rates immediately raise borrowing costs (D_t rises). Furthermore, if higher rates dampen demand for firm products, future expected cash flows decline, reducing equity values (E_t declines). Together, these dynamics lead to a higher debt-to-equity ratio, which increases

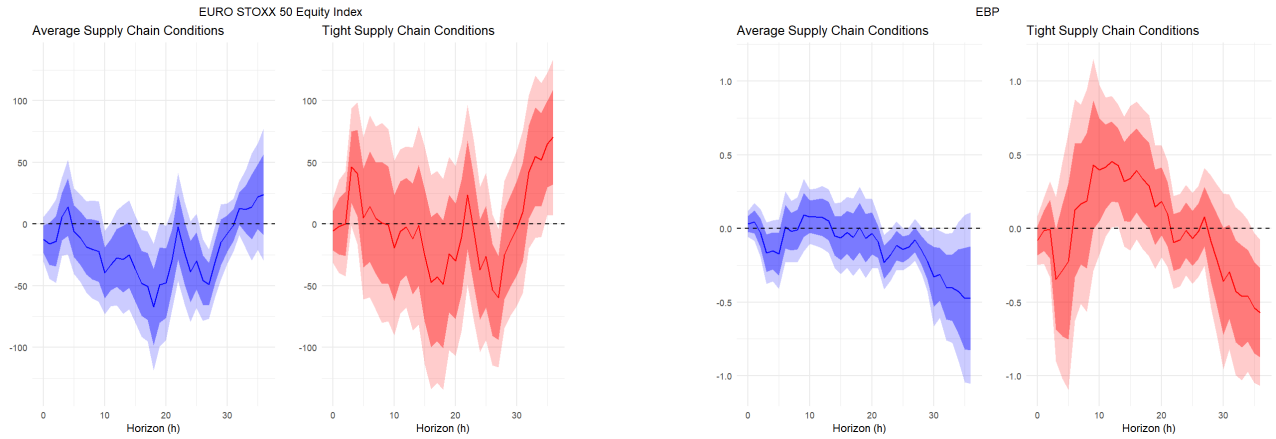
the external finance premium $f\left(\frac{D_t}{E_t}\right)$, making external financing even more expensive.[29]
Laumer and Schaffer (2025) argue that the sensitivity of the external finance premium to changes in the risk-free rate intensifies during periods of supply chain stress. This increased responsiveness can be due to several factors. When supply chains are disrupted, firms face greater uncertainty regarding the availability of inputs or the timing of downstream demand, prompting them to maintain larger inventory buffers or rely more heavily on working capital. As a result, a given increase in interest rates leads to a disproportionately larger increase in borrowing costs (i.e. D_t increases more than usual). Alternatively, supply-side bottlenecks or demand fluctuations can cause firms to revise expected future cash flows downward more sharply (E_t decreases more than usual) in response to higher rates. Lastly, during supply chain disruptions, lenders and investors can place greater emphasis on the financial position of a firm, implying that $f'_{tight} > f'_{average}$. In all these scenarios, the external finance premium $f\left(\frac{D_t}{E_t}\right)$, and consequently the firms interest rate R_t, respond more strongly to changes in R_t^*.

5.2 The Empirical Test

If the proposed hypothesis holds, the cost of external financing should exhibit greater sensitivity to changes in monetary policy during periods of global supply chain disruptions. To assess this, I employ the Laumer and Schaffer (2025) non-linear local projection method, as introduced in the first analysis within the econometric section. Specifically, I estimate the state-dependent IRFs for two financial indicators that reflect external financing costs: the EURO STOXX 50 Equity Index and a European measure of the excess bond premium (EBP).
As in previous analyses, the key parameters of interest are β_0^h and δ_0^h, where $\{\beta_0^h\}_{h=0}^H$ and $\{\beta_0^h + \delta_0^h\}_{h=0}^H$ represent the impulse responses of these financial variables to a monetary policy shock under normal

[29]This simplified explanation, based on Laumer and Schaffer (2025), is sufficient for the purposes of my analysis. Bernanke and Gertler (1995) concisely summarizes the economic rationale behind the credit channel theory as follows: " *According to the credit channel theory, the direct effects of monetary policy on interest rates are amplified by endogenous changes in the external finance premium, which is the difference in cost between funds raised externally (by issuing equity or debt) and funds generated internally (by retaining earnings). The size of the external finance premium reflects imperfections in the credit markets that drive a wedge between the expected return received by lenders and the costs faced by potential borrowers. According to the "credit view," a change in monetary policy that raises or lowers open-market interest rates tends to change the external finance premium in the same direction. Because of this additional effect of policy on the external finance premium, the impact of monetary policy on the cost of borrowing broadly defined and, consequently, on real spending and real activityis magnified. Why should actions taken by the central bank have any effect on the external finance premium in credit markets? In this article we describe two possible linkages. The first of these, the balance sheet channel, stresses the potential impact of changes in monetary policy on borrowers' balance sheets and income statements, including variables such as borrowers' net worth, cash flow and liquid assets. The second linkage, the bank lending channel, focuses more narrowly on the possible effect of monetary policy actions on the supply of loans by depository institutions.*". For a clearer understanding of the credit channel theory of monetary policy, please refer to Bernanke and Gertler (1995), Bernanke (1999), Bean et al. (2002), Gertler and Karadi (2011), Gertler and Karadi (2015).

supply chain conditions and during heightened supply chain stress, respectively. The results are presented in Figure 19.



(a) EURO STOXX 50 (b) EBP

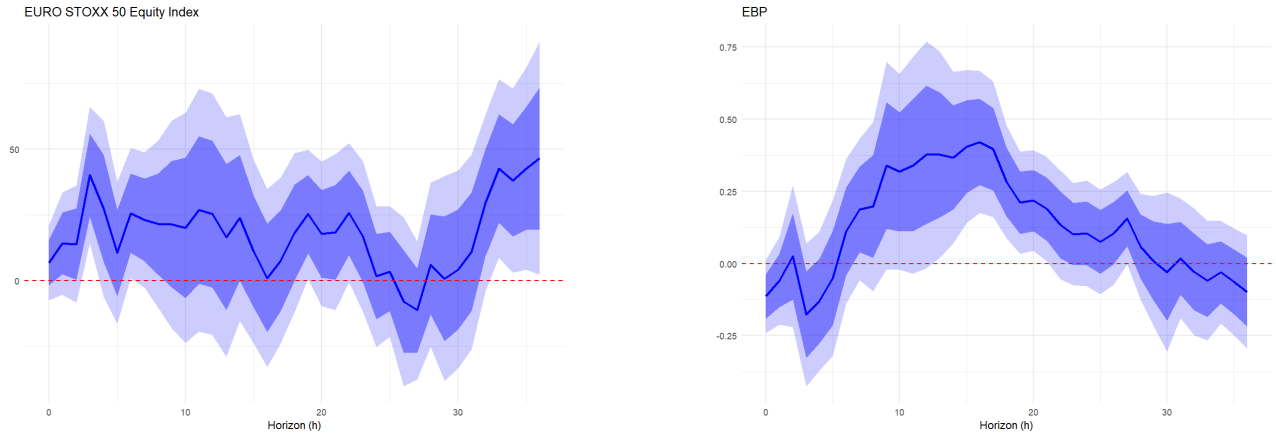
Figure 19: State-dependent impulse responses. Financial variables

The reaction of stock prices is predominantly negative, aligning with standard economic theory. In addition, there is no apparent significant variation between the two states, suggesting that equity values do not exhibit a greater sensitivity to monetary policy during periods of elevated global supply chain pressures.

The excess bond premium (EBP), instead, represents the portion of the interest rate spread between a corporate bond index and a government bond of comparable maturity that is not attributable to default risk. As such, it captures the additional credit costs faced by private firms compared to government securities, solely due to financial frictions. The EBP, therefore, serves as a direct indicator of the external finance premium for large corporations. Thus, investigating its IRFs is particularly relevant for the purpose of my empirical test. Since there is no precise equivalent measure for the European market, I use a closely related variable that reflects a similar economic mechanism.[30] For clarity, I continue to refer to this measure as the EBP. The response of the EBP remains largely neutral when global supply chain pressures are at their average level, with a downward trend emerging at longer horizons. However, under tight supply chain conditions, consistent with the hypothesis of Laumer and Schaffer (2025), the EBP appears to be more sensitive to monetary policy actions. It initially increases, remaining positive until horizon 20, before gradually declining, mirroring the pattern observed under normal conditions.

Figure 20 displays the differential response, namely the $\{\delta_0^h\}_{h=0}^H$ coefficients that capture potential state dependencies. This figure broadly supports the previously noted mixed evidence. Stock prices appear less responsive to monetary policy when global supply chains are under stress, though this difference is rarely statistically significant. In contrast, the IRF of EBP - particularly relevant given the economic mechanism it captures - follows a different pattern. These last state dependent IRFs

[30]See the Data section for more detailed information



(a) EURO STOXX 50 (b) EBP

Figure 20: Differential Responses. Financial variables

indicate increased sensitivity to monetary policy under supply chain stress, with this difference being more statistically significant, especially at medium horizons. This finding corroborates Laumer and Schaffer (2025)'s theory of an intensified role of the credit channel.

Taken together, the evidence is mixed, providing only partial support for the increased sensitivity of the external finance premium when global supply chains are under strain. Notably, this is not inconsistent with the empirical findings presented above, which only partially confirm the amplification mechanism highlighted by Laumer and Schaffer (2025). Therefore, while this analysis does not invalidate the intensification of the credit channel mechanism, it suggests that its role appears to be secondary in the Euro Area and its constituent countries.

6 Conclusion

Using non-linear adaptations of the local projection method of Jordà (2005), this study investigates how the effectiveness of monetary policy is shaped by global supply chain conditions. Overall, I find that stabilizing inflation amid global supply chain pressures becomes slightly easier.

My initial econometric specification, based on Laumer and Schaffer (2025), highlights a weak and rarely statistically significant amplification mechanism for CPI, without substantial heterogeneity between the European countries analyzed, and a modest amplification effect for industrial production. No systematic state-dependent pattern emerges when focusing on the responses of unemployment and retail trade.

Given the European Central Bank's primary mandate of maintaining price stability and the weak statistical significance of the CPI result in this initial analysis, I develop new econometric methods inspired by Hernández et al. (2024) and Jordà et al. (2020). Using a smooth transition autoregression (STAR) model (Granger and Terasvirta, 1993) to explore the state dependence between global supply chain disruptions and monetary policy transmission, where monetary shocks are identified using an instrumental variable strategy, I find an overall modest amplification mechanism for CPI. This second

set of results again reveals slight heterogeneity across the analyzed samples: a modest amplification effect is observed in the Euro Area, France, Spain and Italy; a weaker effect in Denmark; and no state-dependent differences in Germany.

To interpret these findings, I test the hypothesis proposed by Laumer and Schaffer (2025) using data from the Euro Area. According to these two authors, the amplification effect is driven by an intensified credit channel. An empirical test is developed to assess this hypothesis. While the results do not refute the intensification of the credit channel, they suggest its role is secondary in the Euro Area.

Leaving aside the magnitude of the results, focusing on inflation responses, my analysis aligns more closely with Laumer and Schaffer (2025) and Bai et al. (2024), confirming that the sensitivity of inflation to monetary policy increases under global supply chain stress. At the same time, my findings diverge from those of Hernández et al. (2024) and Andriantomanga et al. (2023), who, examining similar questions in emerging and low-income economies, report a significantly dampened monetary policy effect on inflation when global supply chains are under stress. They attribute this to second-round effects on inflation expectations. Taken together, a pattern starts emerging: advanced economies with credible monetary frameworks better anchor inflation expectations and control inflation, while emerging and low-income economies struggle due to weaker monetary frameworks, increasing the risk of inflation expectations de-anchoring.

Finally, shifting the focus to real variable responses, my initial econometric model does not reveal a clear pattern. While a modest amplification mechanism is observed for industrial production, unemployment and retail trade state-dependent differences remain ambiguous or statistically insignificant. Understanding the impact of monetary policy on these real variables amid global supply chain disruptions is crucial for assessing the economic costs of inflation stabilization in similar scenarios. My analysis primarily addresses inflation dynamics, in line with the European Central Bank's primary mandate to maintain price stability, but does not offer a definitive answer on real economic effects. Further research in this direction would be highly valuable.

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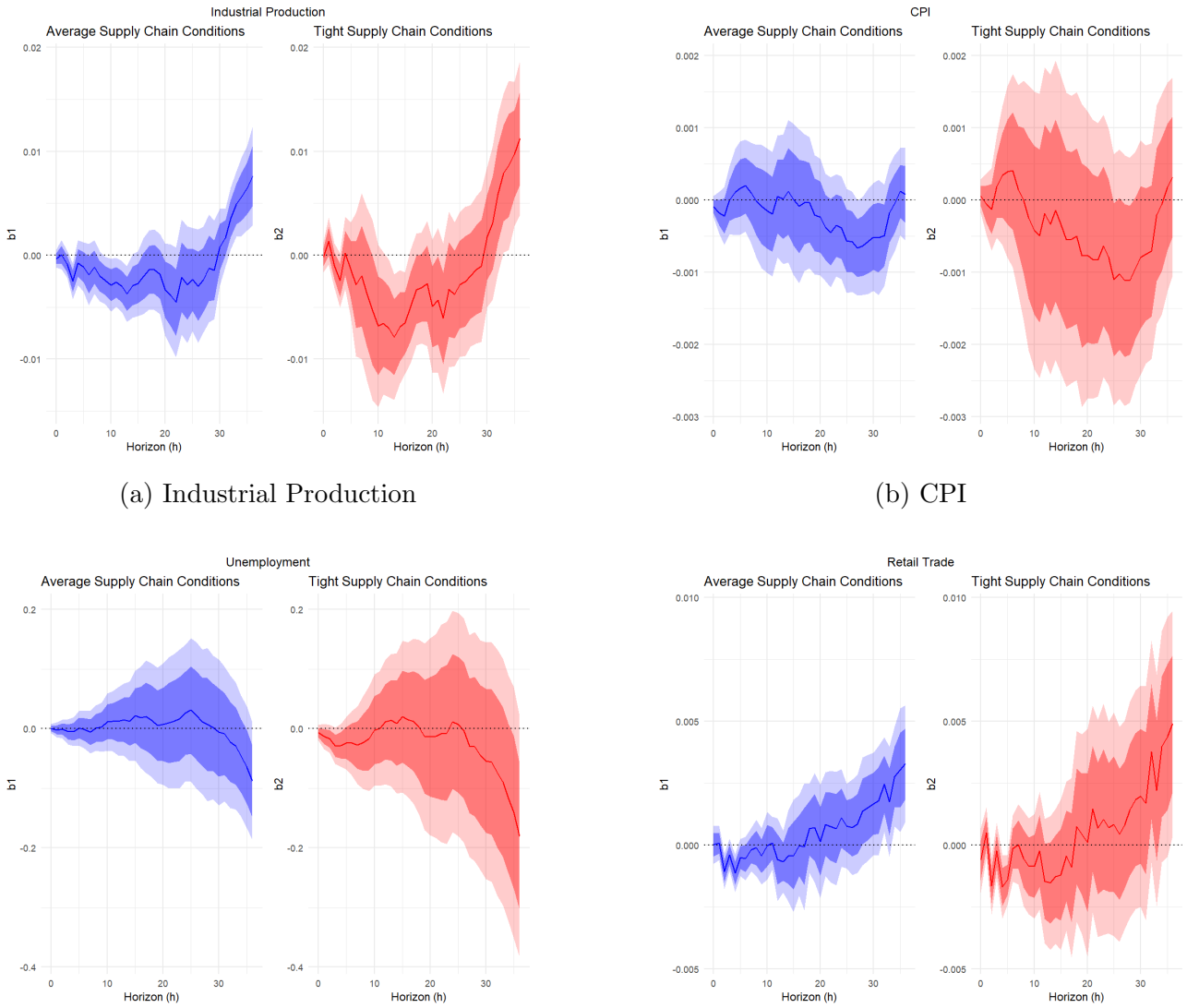
7 Appendix

In this section, I will perform several robustness exercises.

7.1 First Analysis



(a) Industrial Production

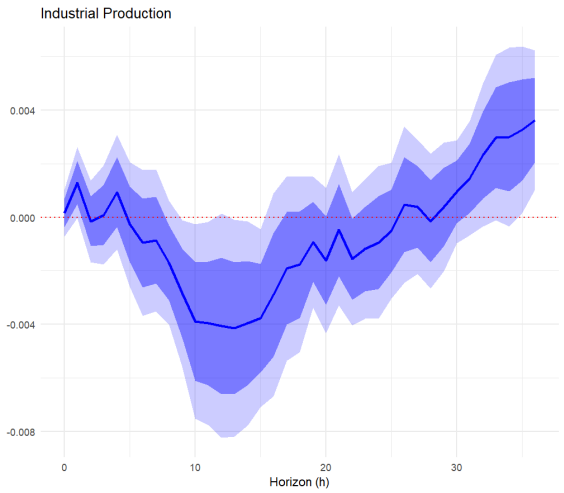
(b) CPI

(c) Unemployment Rate

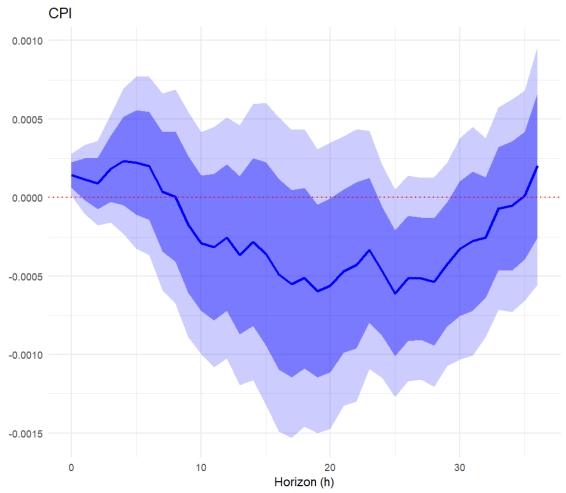
(d) Retail Trade

Figure 21: Euro Area State-dependent IRFs (2019)

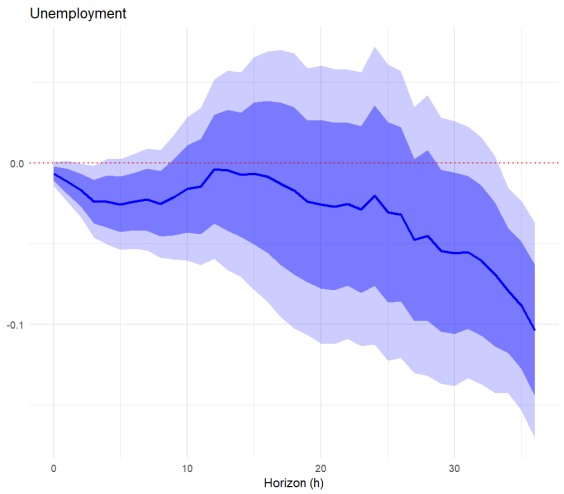
Figure 21 shows the Euro Area state-dependent IRF obtained faithfully applying the Laumer and Schaffer (2025) methodology, that is, also including lags of the other macro variables. Again, the number of lags suggested is set using the Bayesian Information Criterion as spelled out in the econometric section and suggesting 2 lags. However, note that the results are robust to the number of lags



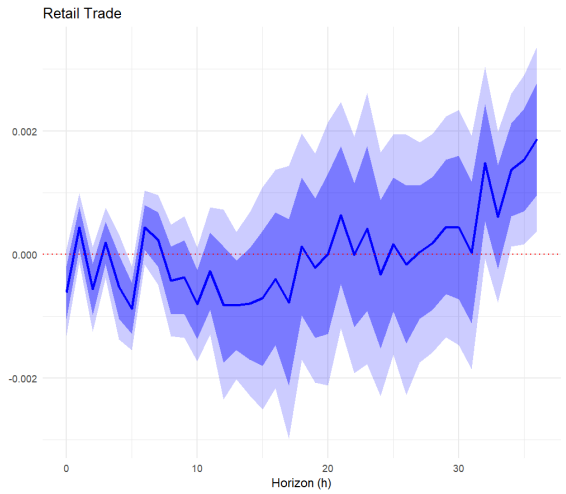
(a) Industrial Production



(b) CPI



(c) Unemployment Rate



(d) Retail Trade

Figure 22: Euro Area Differential Responses (2019)

included. The analysis uses data from January 2001 to December 2019.[31] The results remain fairly robust, and the overall takeaway of a modest or weak amplification mechanism for CPI and industrial production remains unchanged. However, there are some minor differences. The CPI response appears more economically meaningful, deviating from the money neutrality observed in the Euro Area IRF presented in the main text, and exhibiting a weak amplification mechanism. Notably, while a similar CPI response did not emerge at the Euro Area level in the baseline analysis included in the main text, it was consistently observed across nearly all individual country analyses, reinforcing the general takeaway.

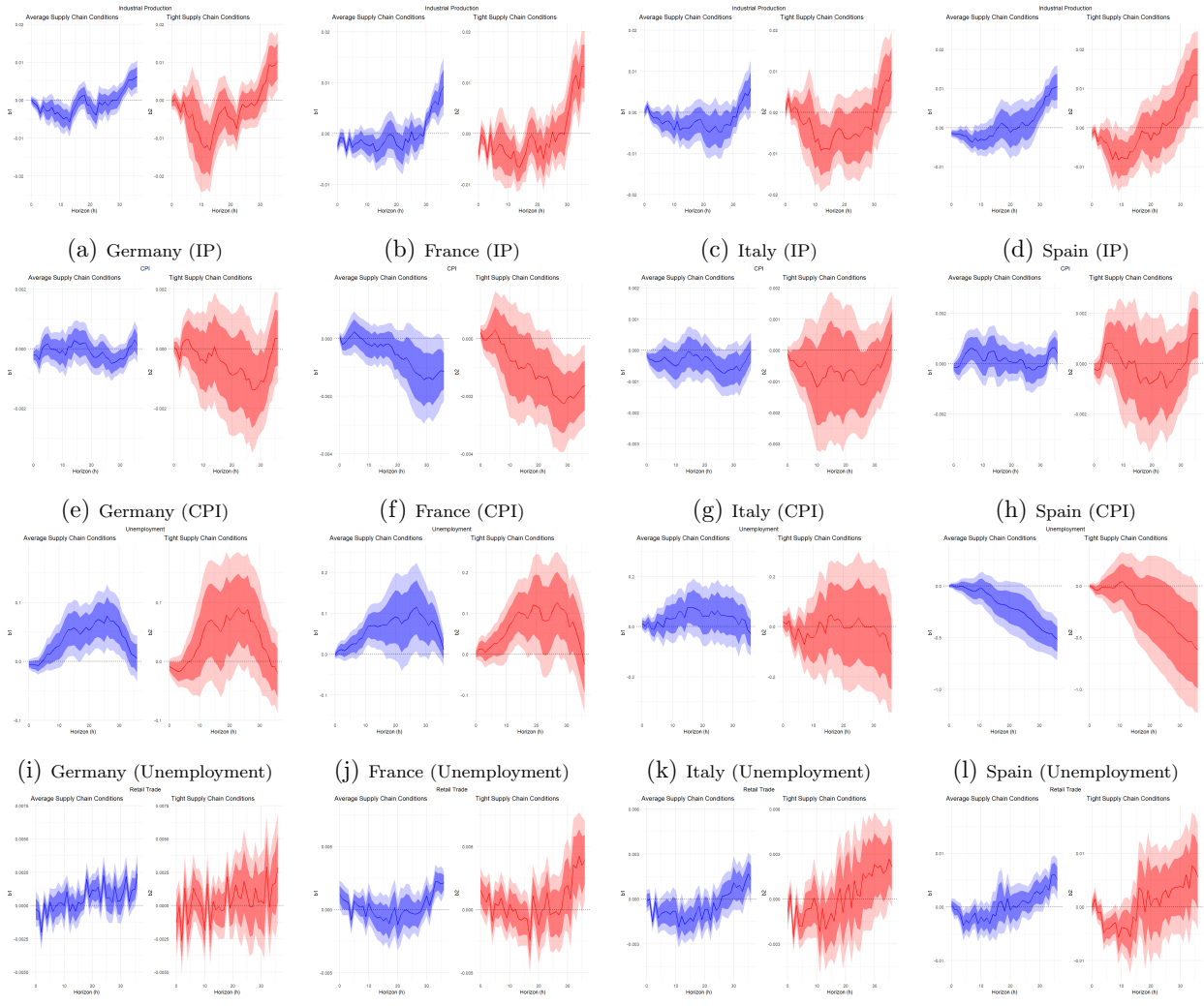
No clear amplification mechanism is observed for retail trade, which appears less economically sig-

[31]For space purposes, I include only the robustness exercises for the baseline analysis, covering observations up to December 2019.

nificant. However, this result was already weakly supported and not confirmed in individual country analyses, so the overall conclusion remains unaffected.

Finally, the unemployment rate exhibits greater money neutrality, with indications of an inverted amplification effect. However, since these responses are not statistically significant, the general takeaway remains unchanged.

To gain additional insights into state dependency in the response, Figure 21 displays the $\{\delta_0^h\}_{h=0}^H$ coefficients. Figure 21 confirms the presence of a weak amplification mechanism for industrial production and CPI. However, at longer horizons, industrial production exhibits the opposite of an amplification effect. For the unemployment rate, no statistically significant differences are observed, with somehow the opposite of an amplification effect. Finally, the response of retail trade display a highly irregular pattern, with signs of the opposite of an amplification mechanism at longer horizons. I now test



(a) Germany (IP) (b) France (IP) (c) Italy (IP) (d) Spain (IP)

(e) Germany (CPI) (f) France (CPI) (g) Italy (CPI) (h) Spain (CPI)

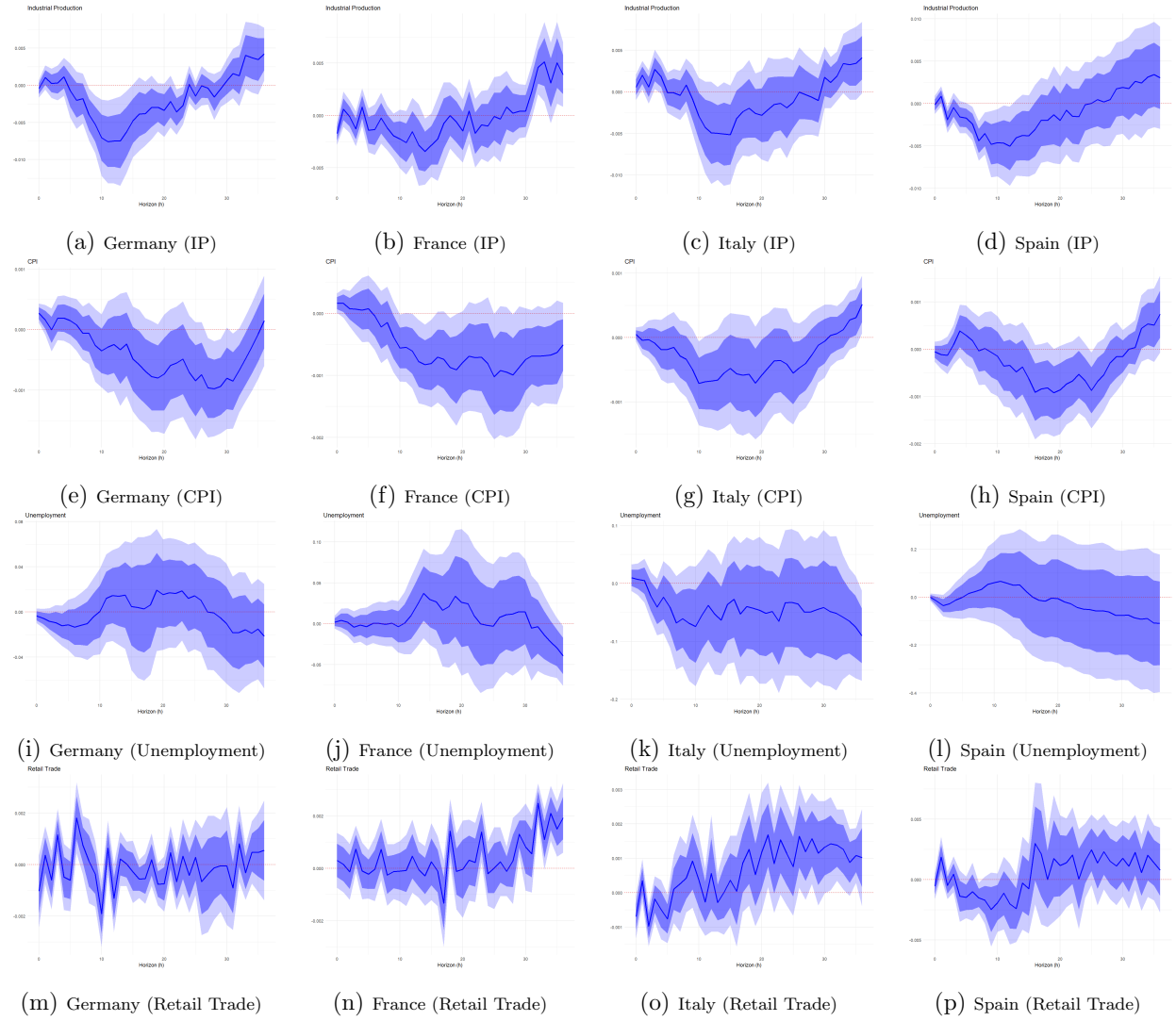
(i) Germany (Unemployment) (j) France (Unemployment) (k) Italy (Unemployment) (l) Spain (Unemployment)

(m) Germany (Retail Trade) (n) France (Retail Trade) (o) Italy (Retail Trade) (p) Spain (Retail Trade)

Figure 23: Euro countries state-dependent impulse responses. Macro variables (2019)

the robustness of the individual countries analyses. Figure 23 displays the individual countries state-dependent IRF. Overall, results are very similar to the baseline analysis included in the main text. CPI seems to be on average more negative with the exception of Spain where a new price puzzle at

shorter horizons emerge (as in the baseline analysis included in the main text). Industrial production is more positive at longer horizons. Unemployment and retail trade response are very similar to the baseline analysis. To get insights on state-dependent effects I inspect Figure 24: the main takeaway



(a) Germany (IP) (b) France (IP) (c) Italy (IP) (d) Spain (IP)

(e) Germany (CPI) (f) France (CPI) (g) Italy (CPI) (h) Spain (CPI)

(i) Germany (Unemployment) (j) France (Unemployment) (k) Italy (Unemployment) (l) Spain (Unemployment)

(m) Germany (Retail Trade) (n) France (Retail Trade) (o) Italy (Retail Trade) (p) Spain (Retail Trade)

Figure 24: Euro countries differential responses. Macro variables (2019)

remains unchanged. No systematic state-dependent differences emerge for retail trade and unemployment, while CPI and industrial production exhibit weak to modest amplification effects. The only notable difference is that, in the baseline analysis, the amplification effect appeared slightly stronger for industrial production, whereas here, the CPI response shows a somewhat greater amplification effect. Finally, at longer horizons, the industrial production response appears to exhibit the opposite of an amplification effect, though not strongly enough to alter the overall conclusion.

7.2 Second Analysis

Figures 25 and 26 show the IRF obtained by setting $\lambda = 14400$ in the Hodrick-Prescott filter. The results are very robust along this dimension.

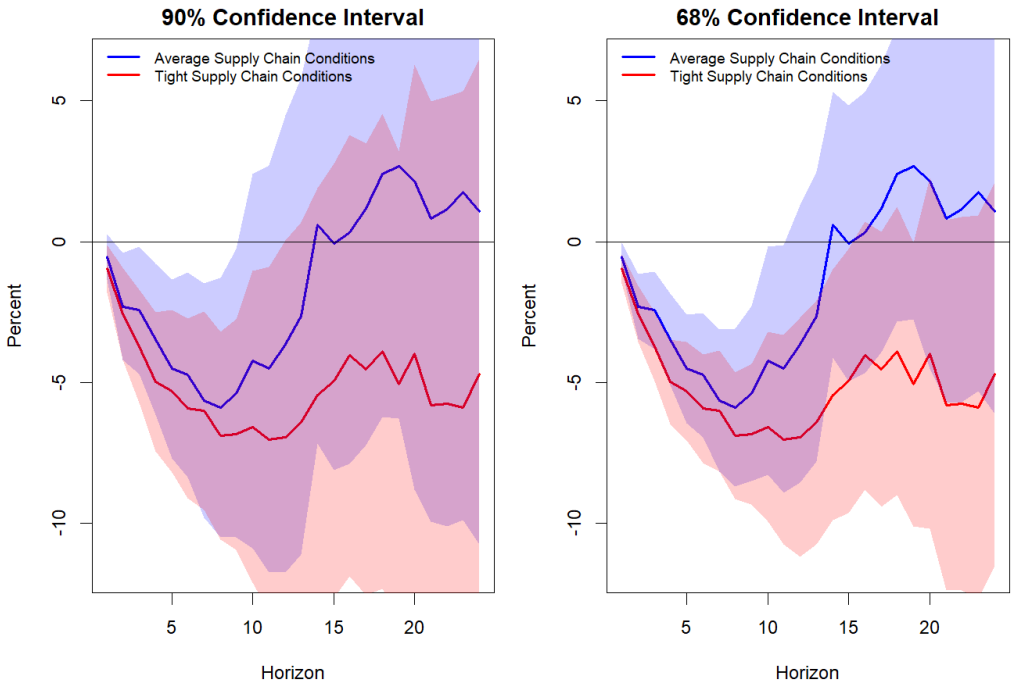
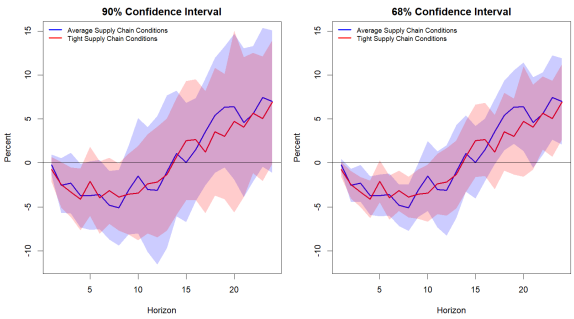
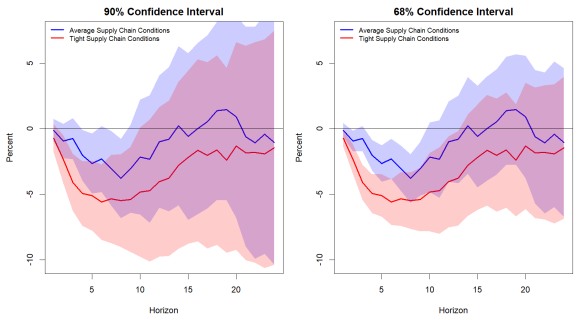


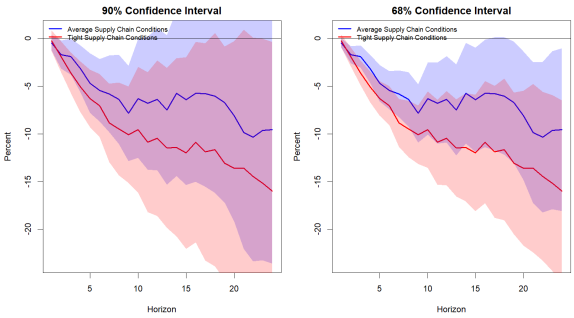
Figure 25: Accumulated Response of Consumer Price Index (log-difference)



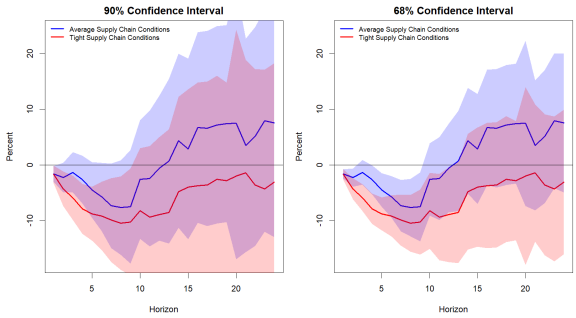
(a) Germany



(b) France



(c) Italy



(d) Spain

Figure 26: Accumulated Response of Consumer Price Index (log-difference)

Figure 27 shows IRF obtained setting $\gamma = 2, 4, 8$ in the logistic function. For space purposes, I only include the IRFs of the Euro Area, nothing changes in the individual countries analyses. They are very robust along this dimension:

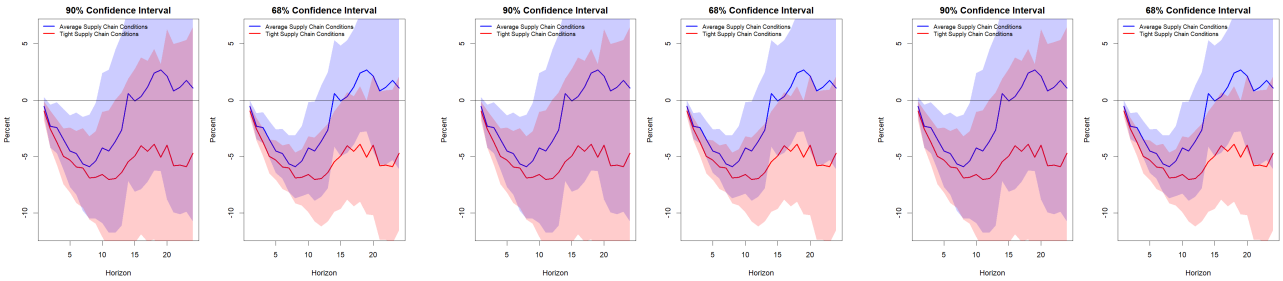


Figure 27: Accumulated Response of Consumer Price Index (log-difference)

Finally, I demonstrate that the results remain highly robust to the number of lags included in the specification. For brevity, I present only the IRFs for the Euro Area, while noting that the individual country analyses also exhibit strong robustness along this dimension. Figure 28 shows the IRF obtained using two lags.

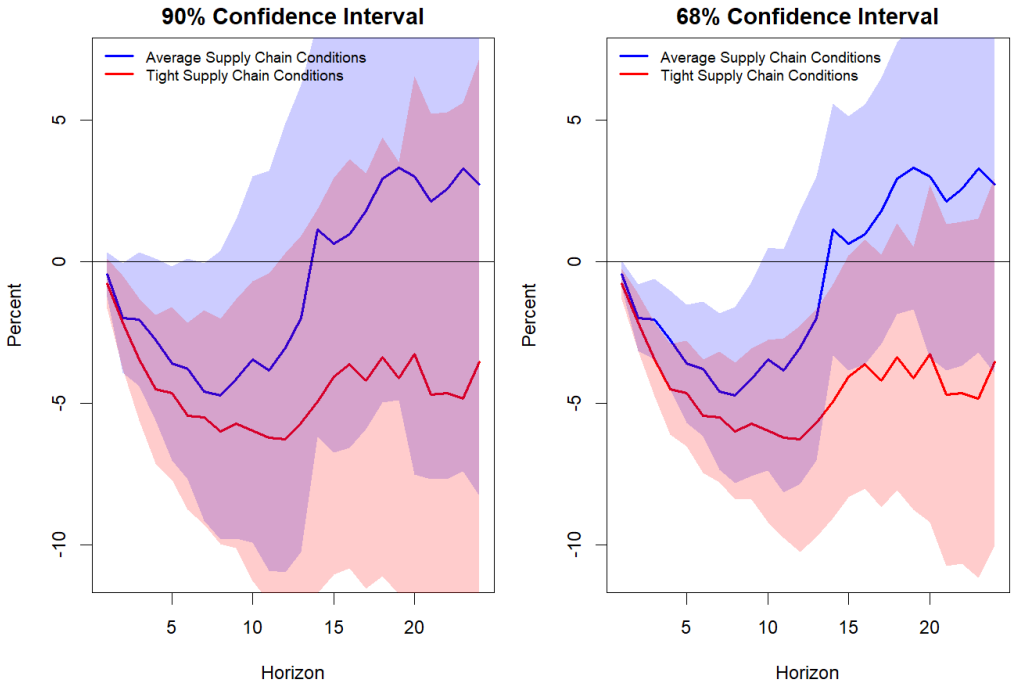


Figure 28: Accumulated Response of Consumer Price Index (log-difference)

7.3 Third Analysis

Robustness checks for the third analysis focusing on Denmark are unnecessary, as the results are virtually identical to those presented in the previous subsection.