# Financial Integration and the Transmission of Monetary Policy in the Euro Area[*]

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**Abstract**

Financial integration determines whether monetary policy transmits across a currency union. Using the euro area—where integration has varied markedly over time—we estimate local projections interacting identified monetary shocks with a continuous measure of integration across member states. Under high integration, monetary policy has strong effects on both consumer prices and output. Under low integration, the same shock has no detectable effect on either. The amplification survives controls for competing state variables, pre-crisis samples, placebo tests, and alternative shock measures. We trace the amplification to three channels: pass-through to government borrowing conditions across member states, cross-border credit supply through integrated banking markets, and portfolio rebalancing through integrated equity markets. The amplification is present in both core and peripheral economies but significantly larger in the periphery.

**Keywords:** Monetary Policy, Financial Integration, Monetary Union, Local Projections

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# 1 Introduction

Does the degree of financial integration shape whether a single monetary policy transmits effectively across a currency union? In a fully integrated area, cross-border capital flows tend to align credit conditions across member states, so that a policy rate change passes through more uniformly; when integration is low, borrowing costs partially decouple from the common stance as local risk premia absorb a larger share of the variation. In the euro area (EA), these are not hypothetical extremes—integration has varied markedly over time, rising steadily in the union's early years, falling sharply during the sovereign debt crisis, and fluctuating even outside crisis episodes. The European Central Bank has long warned that fragmentation impairs transmission, and a large institutional effort—Banking Union, Capital Markets Union—rests on the premise that deeper integration strengthens it. Yet existing empirical work on how fragmentation impairs EA monetary transmission focuses on intermediate financial variables—primarily bank lending rates—without tracing the consequences to aggregate prices and output.

This paper provides that estimate. Using a continuous, quantity-based measure of financial integration from Hoffmann, Kremer and Zaharia (2020)—which tracks intra-EA cross-border asset holdings across money, bond, equity, and banking markets—we interact identified monetary policy shocks with the prevailing level of integration—a union-wide measure that varies over time but not across countries—in a panel local projections framework covering all 19 EA member states from 1999 to 2019. We identify monetary shocks as changes in the two-year German bond rate on ECB Governing Council meeting dates (Altavilla et al., 2019), instrumented by high-frequency surprises from Jarociński and Karadi (2020) to purge information effects. Under high financial integration (75th percentile of the indicator), a one-standard-deviation expansionary shock[1] raises HICP by 0.64% at its peak—about two years after the shock—and real GDP by 0.94% at its peak—about one year. Under low integration (25th percentile), the responses of both variables are small and statistically insignificant throughout the horizon.

The gap between these two regimes is large. At peak horizons, the difference reaches 0.83 percentage points for consumer prices and 0.78 percentage points for output. A one-standard-deviation increase in financial integration[2] shifts the peak HICP response by 0.53 p.p. and the peak GDP response by 0.50 p.p. The two variables respond on different timescales. The output amplification peaks within a year and fades after three years; the price amplification builds gradually, reaching its peak at about two and a half years, and remains significant at longer horizons. This pattern is broadly consistent with standard transmission dynamics: the initial demand stimulus dissipates as supply adjusts, while cumulative price adjustment continues through the Phillips curve. The differential persistence is less pronounced in the pre-2007 subsample, suggesting that crisis-period dynamics contribute to the divergence.

Through which channels is this amplification transmitted? We trace the mechanism in two steps. The first examines the entry point of transmission: the pass-through of the policy rate to sovereign yields across member states. Using intraday yield changes around 268 ECB

[1] Approximately 5.6 basis points on the two-year German bond yield, the standard deviation of the shock series.
[2] This corresponds to an increase of approximately 0.087 units in the quantity-based composite indicator.

Governing Council meetings, the common policy surprise explains 84% of the variation in 2-year sovereign yield changes when integration is high, but only 52% when it is low. The degradation extends across the yield curve—at the 10-year maturity, the share drops from 33% to 12%. Financial integration is thus closely linked to how uniformly the ECB's stance reaches national financial markets.

Uniform pass-through to government yields is not the whole story. The second step decomposes financial integration into four market segments—money, bond, equity, and banking—using price-based sub-indices from Hoffmann, Kremer and Zaharia (2020), and asks which dimensions carry the amplification from financial conditions to final outcomes. The answer is that multiple channels matter, and they map onto different variables. *Banking market integration* has its strongest effect on output, consistent with a credit supply channel in which integrated interbank markets allow a rate cut to lower funding costs uniformly across member states (Cetorelli and Goldberg, 2012; Bruno and Shin, 2015). *Equity market integration* has its strongest effect on prices, consistent with a portfolio rebalancing channel in which cross-border equity holdings propagate asset valuation changes, generating wealth effects that sustain consumer demand (Fratzscher, Lo Duca and Straub, 2016). Bond market integration—the dimension captured by the pass-through exercise—contributes moderately to output amplification but does not drive the price response. No single segment subsumes the composite. The amplification is multi-dimensional.

The mechanism has a clear prediction for cross-country heterogeneity. Peripheral banking systems face higher sovereign risk premia and more severe interbank market segmentation under low integration, while core countries tend to attract safe-haven capital inflows; the gap between integrated and fragmented credit conditions should therefore be widest in the periphery. Consistent with this reasoning, the regime difference reaches 0.99 p.p. for peripheral HICP versus 0.51 p.p. for the core, and 0.86 versus 0.66 p.p. for GDP.

We subject these findings to a battery of robustness and identification tests, organized in three groups. The first addresses identification directly: augmenting the specification with competing state-variable interactions (the business cycle, sovereign risk, and the output gap) individually and jointly following Cloyne, Jordà and Taylor (2023), conducting placebo and permutation tests, and verifying that the monetary shock is distributionally indistinguishable across high- and low-integration quarters. The second tests sensitivity to data and sample choices—a pre-2007 sample excluding all crisis episodes, crisis down-weighting following Lenza and Primiceri (2022), higher-frequency estimation, an extended sample through 2023, and restriction to original EA members. The third considers alternative shock measures and financial integration indicators. The amplification is robust to competing state variables, pre-crisis samples, placebo and permutation exercises, and alternative shocks.[3] These tests support the interpretation that financial integration is not merely a proxy for cyclical conditions or crisis dynamics. Because financial integration varies over time but not across countries, time fixed effects are infeasible; our identification rests on the assumption that, conditional on controls, the remaining variation in integration is orthogonal to other determinants of the

[3]For GDP, the interaction coefficient is robust to each competing control individually at the short-to-medium horizons where the amplification is concentrated, but attenuates in the joint specification that includes all three simultaneously.

macroeconomic response. The tests above provide support for this assumption, though they cannot rule out all potential confounders.

**Contribution.** Our paper makes two contributions. The first is to estimate how financial integration shapes the transmission of monetary policy to *final* macroeconomic outcomes—consumer prices and real output—within a monetary union. Prior empirical work on fragmentation in the EA focuses on intermediate financial variables. Abbassi et al. (2022), Von Borstel, Eickmeier and Krippner (2016), and Hristov, Hülsewig and Wollmershäuser (2014) document that fragmentation impairs the pass-through of policy rates to bank lending conditions, with heterogeneous effects across core and peripheral countries—but they do not trace the consequences to aggregate prices and output. The closest theoretical contributions address related but distinct questions. Wu, Xie and Zhang (2024) develop a two-country New Keynesian model in which financial integration amplifies monetary policy through consumption-switching and financial intermediary channels; their setting, however, features independent monetary policies and flexible exchange rates, where terms-of-trade movements are the primary transmission mechanism—fundamentally different from a currency union with a single policy rate. Bianchi and Coulibaly (2024) show that integration amplifies cross-border monetary policy *spillovers*, a question about how one country's policy affects another, not about how a single central bank's policy transmits within the union it governs. Our paper takes the step from bank rates to aggregate prices and output under a common monetary authority.

Second, the amplification operates through multiple, complementary financial channels—from the uniformity of the interest rate pass-through, through cross-border credit supply and portfolio rebalancing, to final outcomes. The channel decomposition reveals that banking and equity integration are at least as important as bond integration, and that no single market segment accounts for the full effect. These findings are consistent with the theoretical framework of Vari (2020): when the interbank market fragments along national lines—due to counterparty risk that differs across countries—the single policy rate produces different effective funding costs across member states, as peripheral banks must pay a risk premium while core banks accumulate excess liquidity at the deposit facility rate. Extending this framework to incorporate the credit supply and portfolio rebalancing channels documented in the data rationalizes the multi-channel amplification, the core–periphery heterogeneity, and the linear interaction structure of the empirical specification.[4] Our results also speak to the literature on state-dependent monetary policy—Tenreyro and Thwaites (2016) on the zero lower bound, Curdia and Woodford (2010) on credit spreads, Castelnuovo and Pellegrino (2018) on uncertainty—by identifying financial integration as an additional state variable that shapes the effectiveness of transmission. That integration can reverse sharply—gains accumulated over years eroding substantially within quarters (Hobelsberger, Kok Sørensen and Mongelli, 2022)—makes this state variable particularly consequential for central banks.

**Outline.** Section 2 describes the data. Section 3 presents the empirical framework and discusses identification. Section 4 reports the main results—the amplification effect, its

[4] Appendix C embeds Vari's (2020) interbank segmentation in a standard New Keynesian framework, augmented with direct financial channels for credit supply and portfolio rebalancing.

channels, and core–periphery heterogeneity. Section 5 subjects these findings to robustness and identification tests. Section 6 concludes.

## 2 Data

Two ingredients are needed to estimate the amplification effect of financial integration on monetary transmission: a measure of financial integration that varies over time, and a monetary policy shock that does not systematically vary with the state of integration. The analysis is based on a quarterly panel of all 19 EA member states as of 2019, spanning Q1 1999 to Q4 2019. The sample begins with the launch of the euro and ends in 2019 to avoid statistical concerns due to COVID-19.[5] The two outcome variables are HICP and real GDP, both from Eurostat, expressed in base year 2015 and transformed to natural logarithms.[6] HICP is available at monthly frequency and is converted to quarterly format by averaging the three monthly index values within each quarter to match the real GDP frequency. The following two subsections describe the financial integration indicator and the monetary policy shocks in turn.

### 2.1 Financial Integration

A common hallmark of financially integrated markets—where all participants with similar characteristics face the same rules, enjoy equal access, and receive equal treatment—is the prevalence of cross-border trade and the convergence of prices or returns for assets that differ only by their geographical origin, consistent with the law of one price. Cross-border trade and price convergence arise because agents everywhere pursue the best available opportunities and, operating under uniform rules, generate competitive pressures that eliminate arbitrage possibilities and drive price equalization.

Reflecting these features, the literature[7] distinguishes between quantity-based indicators—which focus on cross-border asset holdings—and price-based indicators—which capture yield or return dispersion—to evaluate the state of financial integration in the EA over time. These two types of composite measures offer complementary perspectives: price indices tend to respond quickly to new information, while quantity indices change more gradually. We use the quantity-based composite indicator developed by Hoffmann, Kremer and Zaharia (2020) as our baseline measure of financial integration (FI), which evaluates cross-border financial integration within the EA based on intra-EA asset holdings. To validate the robustness of our findings, we replace our baseline measure with a price-based indicator from Hoffmann, Kremer and Zaharia (2020) and confirm that our main findings hold.[8]

The indicator measures, at the EA aggregate level, the share of assets held cross-border within the union—that is, intra-EA cross-border holdings as a proportion of total EA holdings

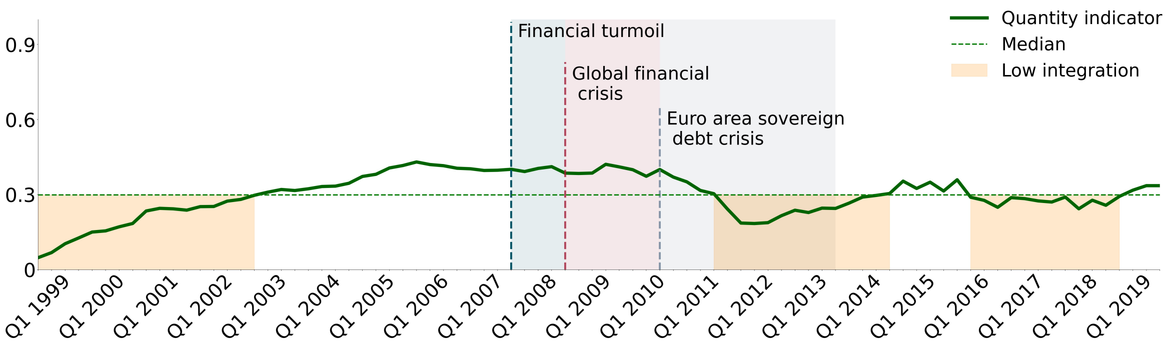
[5] See Lenza and Primiceri (2022) for a detailed discussion of the main statistical challenges posed by the COVID-19 crisis. In Section 5, we extend the sample to Q1 2023 using the variance down-weighting approach of Lenza and Primiceri (2022) and confirm that the main findings are robust.

[6] Real GDP data from Eurostat are seasonally and calendar adjusted, whereas HICP data are not. We verified that our results remain robust when applying seasonal adjustment to HICP; therefore, we present results based on the unadjusted series.

[7] See, for example, Baele et al. (2004) and more recently Hoffmann, Kremer and Zaharia (2020).

[8] See Appendix B.3.4.

**Figure 1:** Quantity-based financial integration composite indicator over time




*Notes:* Quantity-based financial integration composite indicator from Hoffmann, Kremer and Zaharia (2020) is depicted by the dark green line. Median of the indicator is shown as a dotted dark green line. Orange shaded areas indicate periods of low integration (i.e., below sample median). Relevant crisis events are marked based on Hobelsberger, Kok Sørensen and Mongelli (2022).

(cross-border plus domestic)—across four market segments: money, banking, bond, and equity. Within each segment, the relevant assets are: interbank loans (money and banking segments); sovereign bonds held by monetary financial institutions (MFIs) and investment funds (IFs), and corporate bonds held by IFs (bond segment); and equities held by MFIs and IFs (equity segment). These segment shares are then aggregated into a composite indicator using market size-weighted averages. The composite indicator ranges from zero to one, where one signifies complete integration. Further technical details can be found in Hoffmann, Kremer and Zaharia (2020).

The evolution of the resulting quantity-based indicator is depicted in Figure 1, alongside periods of low integration—defined as indicator values below the sample median—and key events identified by Hobelsberger, Kok Sørensen and Mongelli (2022). The indicator began at 0.044 in Q1 1999 and increased steadily in the early years of the monetary union, peaking at 0.423 in Q1 2006. It declined sharply during the global financial crisis and the sovereign debt crisis, reaching a trough of 0.181 in Q2 2012 as capital retrenched from peripheral member states. A partial recovery followed, but the indicator declined again after 2016—outside any major crisis episode—before ending the sample at 0.328 (Q4 2019).

The indicator thus exhibits substantial variation that is not reducible to standard business-cycle fluctuations, suggesting that it captures slow-moving structural characteristics of the EA financial architecture—a characterization we probe formally in Section 5. The indicator rose through the 2001 recession and declined most sharply during the sovereign debt crisis, but it also fluctuated in non-crisis periods—including a notable decline after 2016. Low-integration periods therefore do not exclusively coincide with business-cycle downturns, consistent with the interpretation that the estimated interaction captures variation beyond crisis-related turbulence.[9]

[9]Three robustness checks, presented in Section 5, confirm this result: one augments the main specification with an interaction between monetary shocks and a recession indicator (Appendix B.1.2); one excludes all periods after Q2 2007 (Appendix B.2.1); and one shows that the main results survive competing state-variable interactions (Appendix B.1.1).

We can corroborate that FI is not simply a crisis dummy in disguise by examining how FI variation is distributed across crisis and non-crisis periods. Defining the crisis period as Q3 2007–Q2 2013, only 30.8% of the total sum of squares of the FI indicator (relative to the sample mean) falls within the crisis period, while 69.2% comes from tranquil periods. Within-period variance is virtually identical across regimes: 0.0071 in crisis quarters and 0.0071 outside them, indicating that FI fluctuates just as much during tranquil periods as during stress. For the identifying interaction term—the product of the monetary policy shock and lagged FI—61.3% of the sum of squares falls in crisis quarters, reflecting the larger shock magnitudes during that period, while 38.7% falls outside it. Together, these statistics show that FI provides meaningful variation outside crises, though the majority of the interaction variation (61.3%) falls within the crisis period, where monetary shocks were also largest. The pre-2007 subsample test in Section 5 provides a direct assessment of whether non-crisis variation alone identifies the amplification. Our identification rests on the assumption that, after conditioning on the controls, the remaining variation in financial integration is orthogonal to other determinants of the macroeconomic response to monetary policy—an assumption we probe through our battery of robustness checks (Section 5).

The continuous nature of the FI indicator is key to our empirical strategy: rather than relying on a binary split between low and high-integration periods, we exploit the full time-series variation in integration to assess how monetary transmission depends on the prevailing degree of financial integration. To validate that our findings do not depend on this choice, we conduct three checks: we replace the continuous measure with a binary dummy; test that the linear interaction specification is not rejected by the data; and verify that results scale proportionally across alternative threshold pairs.[10]

### 2.2 Monetary Policy Shocks

Having argued that financial integration provides meaningful time variation that is not reducible to crisis episodes, we now turn to the second ingredient: the monetary policy shock. Following Amberg et al. (2022), we construct the monetary policy shock series in two steps. In the first step, we estimate the following regression on ECB Governing Council meeting dates held between January 1999 and December 2019 (268 meetings, of which 265 have complete yield-change and surprise data; see Appendix A.1), instrumenting the change in the two-year German bond yield with high-frequency monetary surprises, and extract the fitted values $\widehat{\Delta i_d}$:

$$\Delta i_d = \alpha + \varphi \cdot MP_d + \varepsilon_d \tag{1}$$

In the second step, these fitted values are aggregated to quarterly frequency by summing across meeting dates within each quarter, $\widehat{\Delta i_t} = \sum_{d \in t} \widehat{\Delta i_d}$, and enter the panel local projection (Equation 2) as a regressor—that is, Equation 2 is estimated by OLS, not by two-stage least

[10] Appendix B.3.3 (binary indicator), B.1.5 (nonlinearity test), and B.1.6 (threshold comparison) report these results.

squares.[11] In Equation 1, $\Delta i_d$ represents the change in the two-year German bond rate during the monetary event window from Altavilla et al. (2019), and $MP_d$ denotes the monetary shock series identified using the high-frequency identification method from Jarociński and Karadi (2020)[12]. This shock series separates unanticipated monetary policy surprises from informational effects based on the sign of the stock market response to the announcement, addressing concerns about the central bank information channel.[13]

Jarociński and Karadi (2020) employ surprises in financial market prices within a narrow time window around monetary policy announcements, covering both the press statement and press conference. They apply a sign-restriction approach in a two-variable vector autoregression (VAR) to distinguish pure monetary policy shocks from central bank information shocks. These shocks can therefore be interpreted as policy stance changes that are orthogonal to the ECB's communication regarding the economic outlook. These properties make the shocks appropriate for assessing how financial integration modulates monetary transmission.

The results from the estimation of Equation 1 are reported in Appendix A.1. The F-statistic of 79.95 substantially exceeds the conventional threshold of 10 (well above more conservative thresholds in the weak-instruments literature), confirming the strong relevance of the instrument. We report responses to a one-standard-deviation expansionary monetary policy shock (i.e., a decrease in the policy rate), corresponding to approximately 5.6 basis points on the two-year German bond yield ($\sigma_{\widehat{\Delta i_t}} = 0.056$ p.p.).[14]

Table 1 reports summary statistics for the main variables in our panel. The financial integration indicator has a mean of 0.297 and ranges from 0.044 (early years of the monetary union) to 0.423 (pre-crisis peak), with a standard deviation of 0.087. The constructed monetary policy shock has near-zero mean and a standard deviation of 0.056 p.p., consistent with balanced positive and negative surprises.

With these two ingredients in hand—a continuous measure of financial integration and an identified monetary policy shock—we now formalize the empirical specification that tests whether integration amplifies the macroeconomic response to monetary policy.

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[11] A direct 2SLS approach embedding the high-frequency surprises as instruments in the quarterly panel regression is infeasible because the Jarociński and Karadi (2020) shocks, observed only on meeting dates, become weak instruments at the quarterly frequency. Treating the constructed shock as a known regressor in the second step is therefore warranted on two separate grounds. First, the generated-regressor concern is negligible: because the first step is estimated on a large sample of meeting dates with a strong first-stage instrument ($F = 79.95$), the estimation uncertainty in $\widehat{\Delta i_d}$ is small. Second, and independently, Driscoll–Kraay standard errors account for arbitrary cross-sectional and temporal dependence in the panel errors; see Amberg et al. (2022) for a detailed discussion of both arguments.

[12] Our findings hold when we replace the shocks identified via simple ("Poor Man's") sign restrictions with the median rotation shocks that implement the sign restrictions also from Jarociński and Karadi (2020) (Appendix B.3.5).

[13] See, for example, Nakamura and Steinsson, 2018, Cieslak and Schrimpf, 2019, Miranda-Agrippino and Ricco, 2021, and Bauer and Swanson, 2023.

[14] Because $\widehat{\Delta i_t}$ is the quarterly *sum* of instrumented meeting-day changes in the two-year rate, a one-standard-deviation realization represents the cumulative policy-driven shift over an entire quarter—a persistent movement in expected short-term rates over the next two years, rather than a single overnight rate cut. Under the expectations hypothesis, a 5.6-basis-point decline in the two-year yield corresponds roughly to a 5.6-basis-point downward revision in the expected average short rate over the next eight quarters. This persistent nature of the impulse makes the macroeconomic responses larger per basis point than those typically associated with overnight rate surprises.

**Table 1:** Summary Statistics

| | Mean | Std. Dev. | Min | Max | Obs. |
|---|---|---|---|---|---|
| $\Delta$ log HICP (%) | 0.545 | 1.141 | -3.995 | 6.589 | 1577 |
| $\Delta$ log real GDP (%) | 0.594 | 1.508 | -13.133 | 20.688 | 1573 |
| FI indicator | 0.297 | 0.087 | 0.044 | 0.423 | 1596 |
| MP shock ($\widehat{\Delta i_t}$) | 0.003 | 0.056 | -0.212 | 0.226 | 1596 |

*Notes*: $\Delta$ log HICP and $\Delta$ log real GDP are quarterly log-differences (in percentage points). The FI indicator is the quantity-based composite financial integration indicator from Hoffmann, Kremer and Zaharia (2020). The MP shock is the quarterly sum of fitted values from the first-stage regression on ECB Governing Council meeting dates. Sample: Q1 1999 – Q4 2019, 19 EA member states (excluding Croatia). The full panel comprises $19 \times 84 = 1{,}596$ country-quarter observations. The FI indicator and MP shock are available for all country-quarters; HICP and real GDP have 19 and 23 missing observations, respectively, due to data gaps in countries that adopted the euro after 1999.

# 3 Empirical Framework

The intuition outlined in the introduction—that financial integration shapes how uniformly the ECB's policy rate transmits across member states—can be tested using a simple interaction specification. If integration amplifies transmission, then the macroeconomic response to a monetary policy shock should be larger when integration is high. We test this by interacting the shock with the level of financial integration in a local projections framework.

We employ the local projections method of Jordà (2005) to estimate the impulse response functions for the HICP and real GDP with respect to a one-standard-deviation expansionary monetary shock, allowing the response to vary with the degree of financial integration. We estimate the following panel regression for each quarter $h = 0, 1, \ldots, 16$:

$$y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \beta_h \cdot \widehat{\Delta i_t} + \delta_h \cdot \left( \widehat{\Delta i_t} \times FI_{t-1} \right) + \zeta_h \cdot FI_{t-1} + \sum_{k=1}^{K} \gamma_{k,h} X_{i,t-k} + \boldsymbol{\mu}_h' \mathbf{D}_t + \varepsilon_{i,t+h} \quad (2)$$

where $y_{i,t}$ denotes the outcome variable (log of HICP or real GDP) for country $i$ at time $t$, and $\alpha_{i,h}$ are country fixed effects. The term $\widehat{\Delta i_t}$ denotes the series of monetary policy shocks constructed following the two-step procedure explained in Section 2.2; as discussed there, this series enters the regression as a regressor and Equation 2 is estimated by OLS with Driscoll–Kraay standard errors. $FI_{t-1}$ is the lagged measure of financial integration (i.e., the quantity-based composite indicator). The term $\sum_{k=1}^{K} \gamma_{k,h} X_{i,t-k}$ collects controls that include four lags of the outcome variable and one lag of the monetary shock, with the summation running to $K = 4$. The vector $\mathbf{D}_t$ collects three contemporaneous dummies that control for the crisis episodes identified by Hobelsberger, Kok Sørensen and Mongelli (2022): financial turmoil (Q3 2007–Q3 2008), the global financial crisis (Q4 2008–Q2 2010), and the EA sovereign debt crisis (Q3 2010–Q2 2013). We apply a centered moving average over the response coefficients following Jordà and Taylor (2025) to reduce horizon-to-horizon noise.[15] The responses before this adjustment are made available in Appendix A.3. The key parameter is $\delta_h$: a negative $\delta_h$ means that higher financial

[15] Specifically, we use a centered three-quarter uniform moving average (i.e., the smoothed coefficient at horizon $h$ is the simple average of the estimated coefficients at horizons $h-1$, $h$, and $h+1$) implemented via Stata's `tssmooth ma` with `window(1 1 1)`.

integration amplifies the expansionary effect of a rate cut at horizon $h$.

Financial integration enters the specification as a continuous variable rather than a binary indicator. This choice exploits the full time-series variation in the integration measure, avoids making the results dependent on a single threshold, and enables us to estimate the marginal amplification effect at each horizon. The comparison presented in Section 4.2 evaluates the estimated model at the 25th and 75th percentiles of integration, while a robustness exercise using a binary indicator confirms that the amplification effect does not depend on the functional form.[16] Another important feature of this specification is that financial integration indicator varies only over time, not across countries: all member states face the same FI level in a given quarter. This reflects the conceptual nature of the measure—FI captures the degree of integration of the common EA financial area, a property of the monetary union as a whole rather than of individual countries. Country-level measures (such as home bias or bilateral cross-border holdings) would capture different phenomena and raise additional endogeneity concerns. The cross-sectional variation in our panel comes from heterogeneous macroeconomic outcomes across countries, conditional on the common FI level.

Motivated by the fragmentation observed during the sovereign debt crisis, we conduct also a subsample analysis to examine heterogeneity in the effects of monetary policy shocks across two groups of countries: core and periphery. The selection of countries into each group is based on the average level of long-term nominal interest rates for each country between 2010 and 2013. Core countries are defined as those whose 2010–2013 average long-term interest rate was at or below the EA19 aggregate long-term yield from AMECO (the European Commission's macroeconomic database) over the same period, and periphery countries as those with average yields above this threshold.[17] Although the classification is based on crisis-period yields—which could partly reflect the fragmentation dynamics under study—it yields the same classification as Bayoumi and Eichengreen (1992) for all nine countries that appear in both samples (Germany, France, Belgium, and the Netherlands as core; Greece, Ireland, Italy, Portugal, and Spain as periphery).[18]

**Identification.** Equation 2 makes the amplification effect testable, but credible estimation requires that $\delta_h$ is not confounded by omitted variables that co-move with financial integration. The coefficient $\delta_h$ is identified under the assumption that, conditional on the controls, the

[16] See Appendix B.3.3.

[17] Annual data for long-term nominal interest rates were retrieved from AMECO for 2010–2013. For each country, we compute the average yield over this four-year period and compare it to the corresponding EA19 aggregate yield from AMECO (a GDP-weighted average across all 19 member states), averaged over the same period (approximately 3.68%). Core countries (at or below the threshold): Austria, Belgium, Finland, France, Germany, Luxembourg, and the Netherlands. Periphery countries (above the threshold): Cyprus, Estonia, Greece, Ireland, Italy, Latvia, Lithuania, Malta, Portugal, Slovakia, Slovenia, and Spain. Estonia (2011), Latvia (2014), and Lithuania (2015) had not yet adopted the euro during the 2010–2013 classification window; their long-term yields over this period nonetheless reflected market assessments of convergence risk consistent with the periphery classification. The core/periphery classification is robust to small changes in the threshold: shifting the cutoff by $\pm 0.2$ percentage points leaves the composition of both groups unchanged.

[18] Bayoumi and Eichengreen (1992) evaluate the degree of synchronization between supply and demand shocks using data from 1963 to 1989 and a bivariate structural VAR identified via Blanchard-Quah long-run restrictions (AD-AS framework). According to their work, supply-side shocks are highly associated in the core (Germany, France, Belgium, Netherlands, and Denmark) and less so in the periphery (Greece, Ireland, Italy, Portugal, Spain and the UK). Denmark and the UK are included in their sample but are not EA members and therefore do not appear in ours, accounting for the eleven-to-nine country difference between the two samples.

remaining variation in financial integration is orthogonal to other time-varying state variables that independently modulate monetary transmission. Three concerns arise.

First, because FI varies only over time, time fixed effects cannot be included. The absence of time fixed effects is an inherent consequence of using an area-wide measure of financial integration; the indicator measures the interconnectedness of the common financial area and cannot, by construction, vary across countries within the union. The three crisis-period dummies control for direct level effects on outcomes; because they enter additively rather than interacted with the shock, they do not absorb the differential response across integration regimes that $\delta_h$ captures. Any omitted time-varying confounder that co-moves with FI and independently affects the impulse response of prices or output would bias $\delta_h$. The most prominent candidates are the *business-cycle position* (FI may correlate with recessions or expansions), *sovereign risk* (FI fell precisely when spreads widened during the debt crisis), and *institutional reforms* (the OMT announcement in 2012 and progress toward Banking Union restored both integration and ECB credibility simultaneously). The pre-2007 subsample, which predates all non-standard ECB measures and institutional initiatives, provides a direct check on this last concern. These confounders are, however, unlikely to account for our results. The key identifying variation comes instead from two sources: the steady pre-crisis increase in FI (Q1 1999–Q2 2007), over which FI ranges from 0.044 to 0.423— spanning approximately 4.4 standard deviations of the full-sample distribution—that predates the crisis period; and the sharp reversal during the sovereign debt crisis, which provides variation that is distinct from the business cycle. The pre-crisis variation takes the form of a single upward trend rather than fluctuations around a mean, so any other trending variable (e.g., institutional deepening or convergence expectations) could in principle generate a similar interaction pattern—yet a pre-2007 subsample that relies solely on this trend identifies a significant amplification effect, and three placebo EA-aggregate time series with comparable trends fail to replicate it (Section 5). The majority of the interaction-term variation falls within the crisis period (Section 2.1), but crisis down-weighting following Lenza and Primiceri (2022) preserves the amplification at peak horizons (Section 5). Second, if the ECB delivers systematically different shocks during periods of high or low financial integration, the interaction would capture differential shock intensity rather than differential transmission. Third, the linear interaction assumes a constant marginal effect of FI; if the true relationship is nonlinear, the specification may be misspecified.

We address each concern with a dedicated test, reported in Section 5. To assess omitted state variables, we augment the specification with interactions between the monetary shock and three competing candidates—a recession indicator, the sovereign spread (country- and time-varying), and the continuous output gap—individually and jointly in a kitchen-sink specification; the HICP interaction coefficient survives all augmented specifications including the kitchen-sink.[19] A complementary placebo exercise replacing FI with three alternative EA-aggregate time series confirms that only FI produces sustained significance.[20] To assess shock endogeneity, four complementary statistical tests confirm that the monetary shock

[19]For GDP, the interaction coefficient is robust to each competing control individually at the short-to-medium horizons where the amplification is concentrated, but attenuates in the joint specification that includes all three simultaneously. Full results are in Appendix B.1.1.

[20]Appendix B.1.3 reports the full placebo comparison and permutation test results.

is distributionally indistinguishable across high- and low-FI quarters (all $p > 0.4$).[21] To assess the functional form, a quadratic FI interaction following Tenreyro and Thwaites (2016) is insignificant for HICP at all horizons and, for GDP, strengthens rather than weakens the amplification finding; results scale proportionally across alternative threshold pairs.[22] Section 5 reports additional checks—including higher-frequency estimation, an extended sample through 2023 with COVID handling, and alternative shock series—that confirm the amplification.

Two further concerns relate to the specification itself. First, the two-step shock construction (daily first stage, quarterly second stage) creates a generated regressor. Driscoll–Kraay standard errors account for serial and cross-sectional dependence but not first-stage estimation uncertainty. The strong first stage ($F = 79.95$)[23] addresses the attenuation-bias concern: weak instruments would shrink the estimated effects toward zero, but the high $F$-statistic rules this out. A complementary bootstrap procedure (500 replications) re-sampling the first-stage meeting-level regression with replacement and re-estimating the full two-step procedure confirms that the reported Driscoll–Kraay standard errors are conservative, so first-stage estimation uncertainty does not materially inflate inference uncertainty.[24] Second, financial integration enters with one lag, mitigating—though not fully resolving—the possibility that monetary policy itself affects integration contemporaneously (e.g., through QE compressing spreads). The quantity-based nature of the FI indicator nevertheless limits the scope for contemporaneous contamination: cross-border asset holdings are inherently slow-moving and unlikely to respond meaningfully within the same quarter as a policy surprise.

# 4 Results

We present the results in four steps. First, we document the headline finding: financial integration amplifies the response of both prices and output to monetary policy shocks (Section 4.1). Second, we compare impulse responses at low versus high integration levels, showing that monetary transmission is statistically and economically insignificant when integration is low (Section 4.2). Third, we ask whether the amplification differs across core and peripheral member states (Section 4.3). Fourth, we investigate the *channels* through which integration amplifies transmission, first by examining the interest rate pass-through and then by decomposing the composite indicator to identify which financial market segments drive which macroeconomic outcomes (Section 4.4).

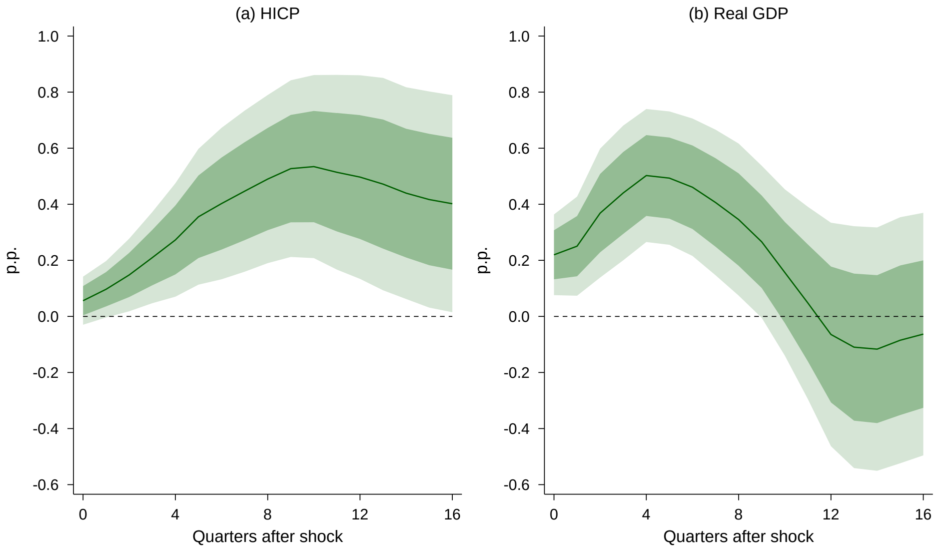
[21] Appendix B.1.4 provides descriptive statistics, formal test results, and kernel density estimates.

[22] Appendix B.1.5 reports the quadratic coefficients and implied IRFs; Appendix B.1.6 provides the threshold comparison.

[23] See Appendix A.1 for the full first-stage results.

[24] The bootstrap yields standard errors that are, if anything, smaller than the Driscoll–Kraay estimates: the average bootstrap-to-DK SE ratio is 0.86 for the interaction coefficient on HICP and 0.58 for GDP. Together, the high $F$-statistic (ruling out attenuation bias) and the bootstrap ratios below unity (ruling out SE inflation) indicate that the generated-regressor issue is not consequential for our results.

**Figure 2:** Amplification effect of financial integration on HICP and real GDP responses




*Notes:* Change in the impulse response (p.p.) associated with a one-standard-deviation increase in financial integration ($\approx$ 0.087), for a one-standard-deviation expansionary monetary policy shock ($\approx$ 5.6 basis points). Shaded areas represent 68% and 90% confidence intervals based on Driscoll-Kraay standard errors.

### 4.1 Amplification Effect of Financial Integration

Because financial integration enters Equation 2 through a linear interaction with the monetary policy shock, the amplification of the impulse response is fully characterized by the interaction coefficient $\delta_h$. The impulse response at horizon $h$ to a shock of size $s$, evaluated at a level of financial integration $FI$, is $(\beta_h + \delta_h \cdot FI) \cdot s$. A one-standard-deviation increase in financial integration ($\sigma_{FI} \approx 0.087$) therefore shifts the impulse response by $\delta_h \cdot \sigma_{FI} \cdot s$ at each horizon, regardless of the starting level of $FI$ (by construction of the linear interaction specification).

Figure 2 reports this amplification effect—the change in the impulse response associated with a one-standard-deviation increase in financial integration—for a one-standard-deviation expansionary monetary policy shock.[25]

For HICP, the amplification effect rises sharply, peaking at approximately 0.53 p.p. two and a half years post-shock and remaining statistically significant (at the 90% confidence level) at longer horizons, indicating a sustained amplification of price responses as financial integration increases. For real GDP, the amplification effect peaks at approximately 0.50 p.p. one year after the shock and fades thereafter, becoming statistically insignificant after about three years. This pattern suggests that greater financial integration provides a persistent boost to prices but only a temporary increase in output. We note that the HICP amplification effect is robust to all identification tests presented in Section 5, including competing state-variable controls in a kitchen-sink specification (a specification that includes all competing controls simultaneously). For GDP, the interaction coefficient is robust to each competing control

[25] Formally, the plotted quantity at horizon $h$ is $-\hat{\delta}_h \cdot \sigma_{FI} \cdot \sigma_{\widehat{\Delta i}}$, where the negative sign converts the shock to an expansionary surprise. This is the derivative of the impulse response with respect to $FI$, scaled by $\sigma_{FI}$ and evaluated at a one-standard-deviation shock.

individually but attenuates in the joint specification (see Section 5 for details).

The differential persistence across prices and output is consistent with standard transmission dynamics: output adjusts relatively quickly as supply responds to the initial demand stimulus, while prices adjust gradually through the Phillips curve. Under high financial integration, the coordinated pass-through of the policy rate generates a uniform demand stimulus across member states—uniform in the sense that all countries face similar borrowing conditions, not necessarily that the stimulus is large in magnitude—producing a large aggregate output gap that then feeds into sustained price adjustment. Under low integration, the fragmented pass-through limits the aggregate demand impulse, attenuating both the output and price response—but the effect on cumulative price adjustment is proportionally larger because the persistent component of the stimulus is what matters most for prices. We treat this asymmetry as descriptive rather than structural, and note that it is less pronounced in the pre-2007 subsample, suggesting that the crisis period—when fragmentation was most severe—contributes to the estimated divergence between price and output dynamics.[26][27]

### 4.2 Monetary Policy Transmission Across Levels of Financial Integration

Figure 3 reports the impulse responses of HICP and real GDP to a one-standard-deviation expansionary monetary policy shock, evaluated at the 25th percentile (approximately 0.242) and 75th percentile (approximately 0.377) of financial integration. The coefficients obtained from estimating Equation 2 are reported in Appendix A.2.[28]

For HICP, the response to an unexpected decrease in the policy rate under low financial integration (25th percentile) is negligible—around 0.04%—and remains statistically and economically insignificant throughout the horizon. In contrast, under high financial integration (75th percentile), the initial response rises to 0.13%, peaking at 0.64% around two years after the shock and remaining statistically significant thereafter. The difference in price responses between the two regimes reaches 0.83 p.p. at the peak—at a horizon where the low-FI response has turned to approximately −0.19%, widening the gap well beyond the initial impact values—indicating that monetary policy transmission to prices strengthens markedly with higher financial integration.

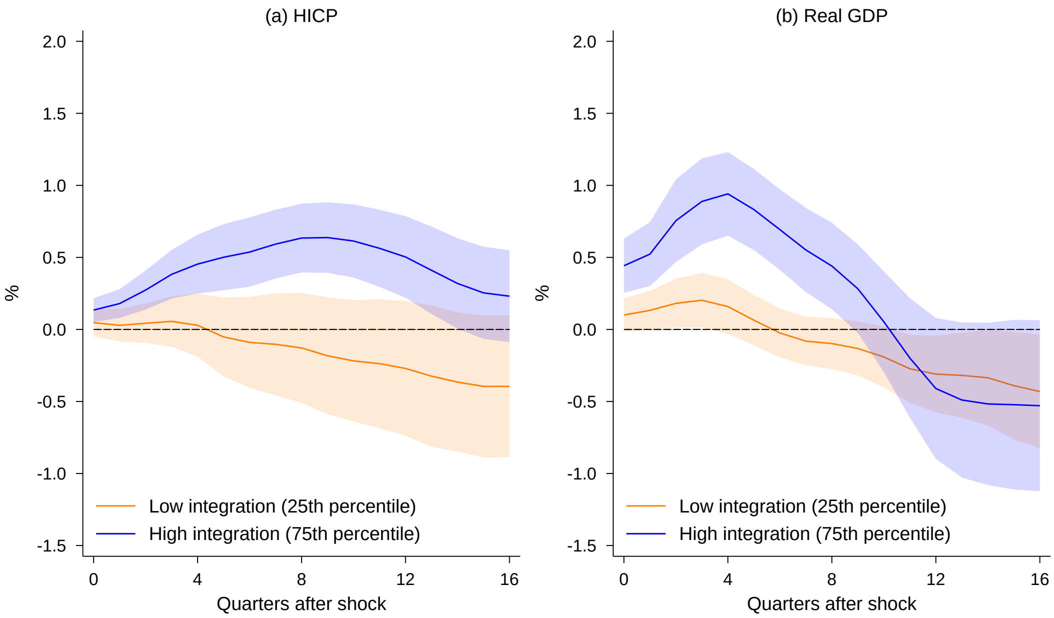
For real GDP, a qualitatively similar pattern emerges. Under low financial integration, the output response is initially small and statistically insignificant, while under high integration it is considerably stronger—starting at 0.44% and peaking at 0.94% one year post-shock. The difference in output responses across regimes reaches 0.78 p.p. at the peak (reflecting a low-FI response of approximately 0.16% at that horizon), slightly smaller than the 0.83 p.p. HICP

[26]Appendix B.2.1 reports the pre-2007 subsample results.

[27]These amplification effects are internally consistent with the two-regime comparison presented in the next subsection: the peak difference between the 75th and 25th percentile regimes (0.83 p.p. for HICP) corresponds to a change in $FI$ of $0.377 - 0.242 = 0.135$, while the amplification effect reported here corresponds to a change of $\sigma_{FI} \approx 0.087$. Under the linear interaction structure, the ratio of peak differences should equal the ratio of FI gaps. Consistently, the ratio of HICP peak differences ($0.83/0.53 \approx 1.57$) closely matches the ratio of FI gaps ($0.135/0.087 \approx 1.55$).

[28]These coefficients represent the values before the following adjustments: (i) rescaling the monetary shock to a one-standard-deviation shock; (ii) changing the sign to an expansionary shock; (iii) scaling by the level of financial integration; and (iv) applying the centered moving average following Jordà and Taylor (2025) to ensure smoothness.

**Figure 3:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Impulse responses are shown in percentage terms to a one-standard-deviation expansionary monetary policy shock, corresponding to approximately 5.6 basis points on the two-year German bond yield. Results are reported for low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

peak difference. However, the GDP response is notable for its earlier peak timing (quarter four versus quarter nine for HICP). Under low financial integration, the point estimates occasionally turn negative at some horizons—implying, counterintuitively, that an expansionary shock could lower prices or output. However, these negative values are statistically insignificant throughout, reflecting estimation imprecision rather than a genuine economic effect.[29]

Taken together, these results indicate a substantial difference between low- and high-integration regimes: at low levels of financial integration, the estimated responses are statistically and economically insignificant throughout the horizon, while under high integration, transmission is strong and precisely estimated. This contrast raises a natural follow-up: does the amplification operate symmetrically across member states, or is it concentrated in countries that were most affected by fragmentation?

**Plausibility of the response magnitudes.** Our average GDP response—evaluated at the sample mean of financial integration ($\approx 0.297$)—peaks at 0.48% about one year after a one-standard-deviation shock of approximately 5.6 basis points on the two-year German bond yield.

[29]Because the impulse response is $(\beta_h + \delta_h \cdot FI) \cdot s$, the intercept $\beta_h$ (corresponding to $FI = 0$) lies outside the support of the data and has no standalone economic interpretation; the relevant comparison is between plausible levels of integration, such as the 25th and 75th percentiles reported here.

This is broadly consistent with existing EA estimates.[30] Three features of our setup inform the comparison. First, because our shock is the quarterly sum of instrumented daily changes in the two-year rate, a one-standard-deviation realization represents the cumulative policy-driven shift over an entire quarter—a persistent impulse more powerful per basis point than a single-meeting surprise. Second, the conditional responses at the 75th percentile are larger than the unconditional average by construction: this amplification is the paper's central finding, and the relevant comparison with the literature is the average response. Third, our estimates are most informative for shocks of the size that actually occur in the data; linear extrapolation to a conventional 25-basis-point shock would overstate the likely macroeconomic effect.

### 4.3 Subsample Analysis: Core versus Periphery

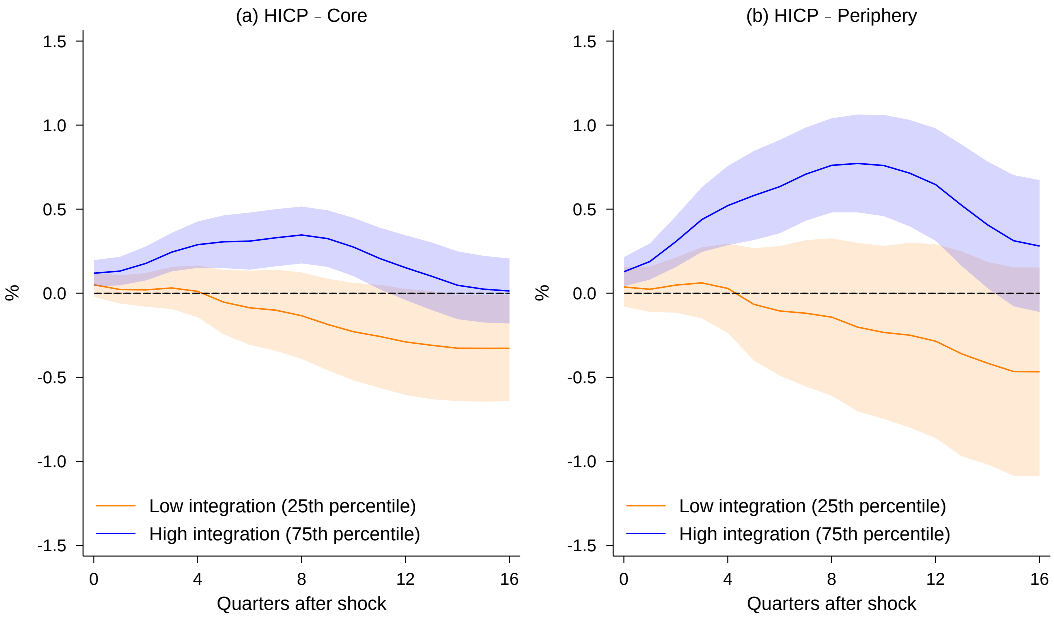
The sovereign debt crisis triggered a clear fragmentation within the EA, differentiating core member states—characterized by lower sovereign risk premia and experiencing capital inflows—from peripheral member states, which experienced significant capital outflows. We examine this heterogeneity by estimating impulse responses separately for core and periphery groups, shown in Figures 4 and 5.

For HICP, peripheral member states exhibit stronger and more persistent responses to the monetary policy shock under high financial integration. The inflation response in the periphery peaks at 0.77% approximately two years after the shock, while the core response peaks at 0.34% exactly two years post-shock and declines more rapidly. The difference in price responses between low and high integration reaches 0.99 p.p. two and a half years after the shock for the periphery and 0.51 p.p. at the same horizon for the core.

Real GDP responses reveal a comparable pattern. Peripheral member states show stronger amplification, with output peaking at 1.07% in the high-integration regime one year after the shock, compared with 0.76% for the core. The difference between low- and high-integration regimes at the peak reaches 0.86 p.p. and 0.66 p.p. one year post-shock for the periphery and core, respectively. Overall, these findings indicate that peripheral member states experience larger amplification effects—more persistent for prices and more pronounced at the peak for output—under high financial integration, whereas effects in the core are more transitory. We note that these subsample estimates are descriptive; we have not formally tested whether the amplification differs significantly across subgroups, and the asymmetric subsample sizes (7 core versus 12 periphery countries) limit the power of such a test.

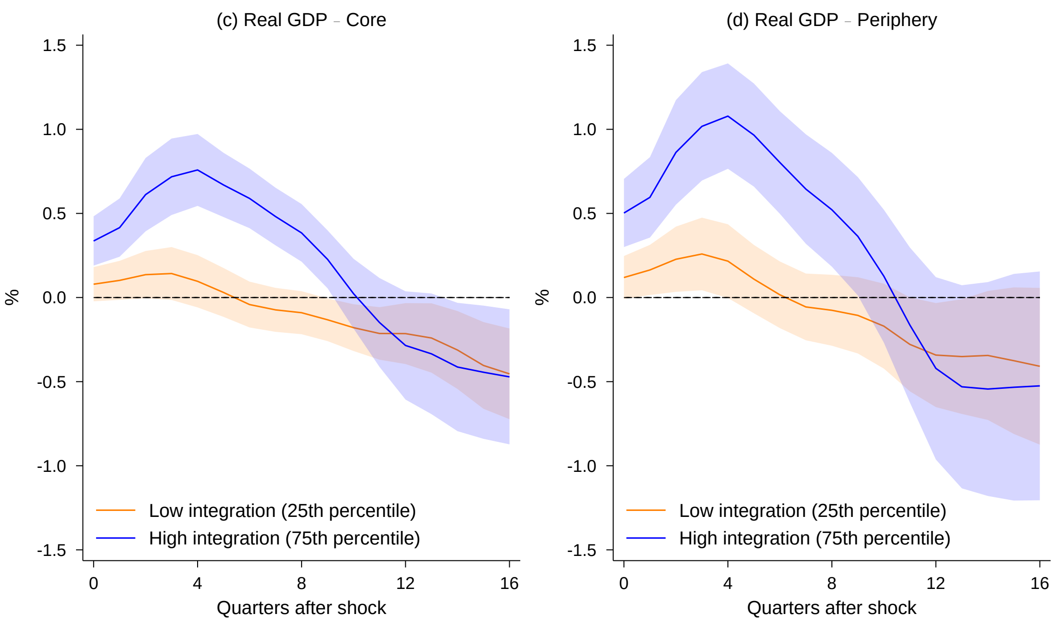
[30] The most directly comparable benchmark is Bauer and Swanson (2023), who also normalize to the two-year yield. Rescaling their peak industrial production response proportionally to our shock size ($\times$ 5.6/25) yields 1.43% (the underlying peak IP response to a 25 bp shock is approximately 6.4%; this US benchmark assumes the EA two-year rate shock is broadly comparable), which exceeds our average GDP response, as expected given that IP is more cyclically sensitive. Even our conditional response at the 75th percentile (0.94%) remains below this benchmark. For the EA, Jarociński and Karadi (2020) report smaller effects, unsurprisingly given five differences in setup: they use monthly rather than quarterly data; a shorter-maturity instrument; a smaller shock (3.8 basis points versus our 5.6 basis points); a shorter sample; and they use the GDP deflator rather than HICP to measure prices. Corsetti, Duarte and Mann (2022) document substantial heterogeneity in country-level responses broadly consistent with ours, and Buda et al. (2025) show that temporal aggregation can shift the timing of effects while keeping magnitudes in a similar range.

**Figure 4:** Responses of HICP to a monetary policy shock under different levels of financial integration for core and periphery




*Notes:* Impulse responses are shown in percentage terms to a one-standard-deviation expansionary monetary policy shock, corresponding to approximately 5.6 basis points on the two-year German bond yield. Results are reported for low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

**Figure 5:** Responses of real GDP to a monetary policy shock under different levels of financial integration for core and periphery




*Notes:* Impulse responses are shown in percentage terms to a one-standard-deviation expansionary monetary policy shock, corresponding to approximately 5.6 basis points on the two-year German bond yield. Results are reported for low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

## 4.4 Mechanism

Sections 4.1–4.3 have documented *that* financial integration amplifies monetary transmission, and that the effect is present in both core and periphery but more pronounced in the latter. We now turn to the question of *how*. Through which channels does financial integration amplify monetary transmission? We begin by examining the most direct channel—the uniformity of interest rate pass-through—before decomposing the composite FI indicator into its constituent market segments to characterize which financial channels drive the amplification of macroeconomic outcomes.

### 4.4.1 Interest Rate Pass-Through

We first examine whether financial integration is associated with the uniformity of sovereign yield responses to ECB policy announcements—specifically, whether the degree to which the common policy surprise passes through to bond yields across member states depends on the level of financial integration.

Using the EA-MPD event-study database from Altavilla et al. (2019), we measure the change in sovereign yields for Germany, France, Italy, and Spain at the 2-year maturity within a narrow window around each ECB Governing Council meeting.[31] We then plot each country's yield change against the change in the 2-year overnight indexed swap (OIS) rate—a pure measure of the common ECB policy surprise, free of country-specific sovereign risk premia—separately for meetings occurring in quarters of high and low financial integration (above and below the sample median of the FI indicator).[32] The pooled $R^2$ from regressing all countries' yield changes on the OIS rate change serves as our summary measure of pass-through uniformity.

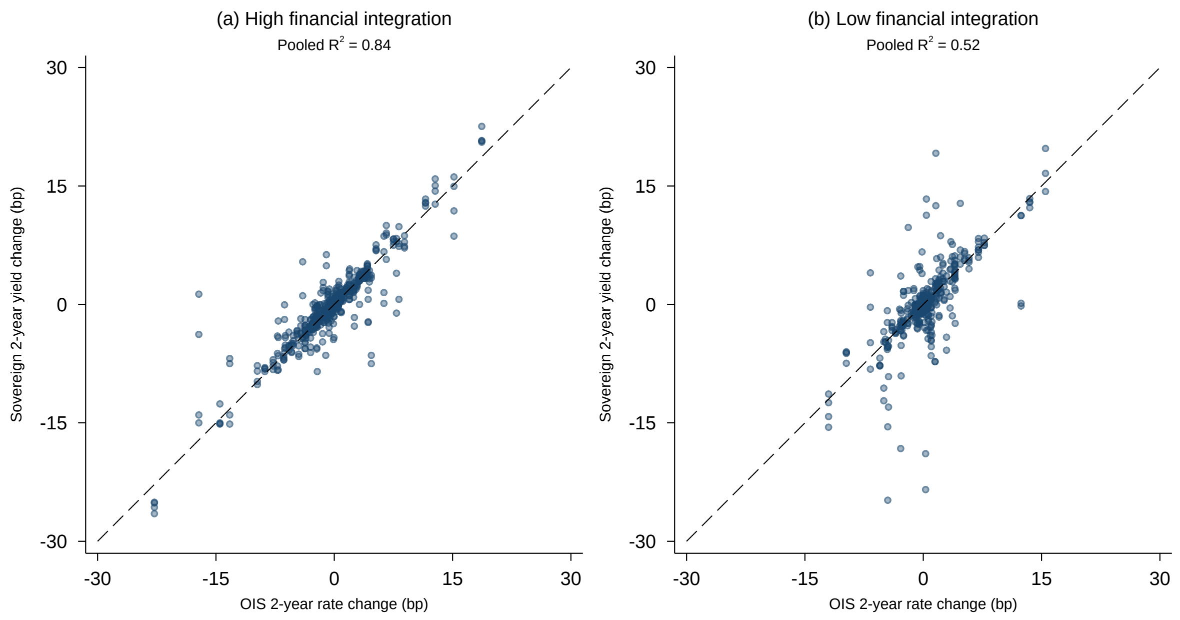
Figure 6 reports the results for the 2-year maturity. Under high financial integration (panel a), sovereign yields from all four countries track the OIS rate closely: pooling across countries, the common policy surprise explains 84% of the variation in 2-year yield changes. Under low financial integration (panel b), the scatter widens considerably, and the pooled $R^2$ drops to 52%. Figure 7 shows that this pattern extends across the yield curve: at the 5-year maturity, the pooled $R^2$ is 70% under high FI and 33% under low FI; at 10 years, 33% and 12% respectively.[33] Importantly, this pattern is not driven by sovereign risk dynamics during the debt crisis: restricting the sample to the pre-crisis period (January 1999 to June 2007), when sovereign risk differentials across member states were negligible, preserves the finding—the $R^2$ remains high

[31]The EA-MPD "Monetary Event Window" spans from just before the press release to just after the press conference, isolating the financial market reaction to the policy decision. We use 268 meetings from January 1999 to December 2019. All yield changes are in basis points. Germany, France, Italy, and Spain are the only EA countries for which intraday sovereign yield changes are available in the public EA-MPD database. These four economies are also the largest in the EA, accounting for approximately three quarters of EA GDP.

[32]Financial integration in the mechanism analysis enters contemporaneously (the FI value of the quarter in which the meeting occurs) rather than with the one-quarter lag used in the main specification. This is appropriate because the mechanism operates within the same quarter: a meeting occurring in a high-FI quarter should exhibit uniform pass-through immediately. The results are robust to using lagged FI instead.

[33]The pass-through *coefficient* (the slope) remains broadly stable across regimes, consistent with the ECB shock remaining a relevant common factor. What changes is the *share of variance* it explains—i.e., the $R^2$—because under low financial integration, country-specific factors generate substantial idiosyncratic yield movements that dilute the common policy signal. We verify this formally in Appendix A.5: an interaction regression of sovereign yield changes on the OIS rate interacted with FI produces insignificant interaction terms at all maturities.

**Figure 6:** Pass-through of ECB policy shock to sovereign 2-year yields




*Notes:* Each point represents one ECB Governing Council meeting (268 meetings, January 1999 – December 2019). The horizontal axis is the change in the 2-year OIS rate (basis points) during the monetary event window from Altavilla et al. (2019); the vertical axis is the corresponding change in the sovereign 2-year yield for Germany, France, Italy, or Spain. The dashed line is the 45-degree line (perfect pass-through). $R^2$ values are from pooled OLS regressions of all countries' yield changes on the OIS rate change, with heteroskedasticity-robust standard errors. High FI: meetings in quarters where the financial integration indicator exceeds the sample median; Low FI: at or below the median.

under high FI and deteriorates markedly under low FI at longer maturities.[34]

The degradation is striking both across regimes and across maturities: the common ECB policy surprise explains 84% of 2-year yield variation under high financial integration but only 12% of 10-year yield variation under low integration—a sevenfold difference that reflects both the regime gap and the maturity gradient. This gradient is consistent with the mechanism: at longer maturities, sovereign-specific risk factors (credit risk, liquidity, fiscal expectations) play a larger role, and financial fragmentation allows these factors to dominate the common policy signal. Notably, the absolute drop in $R^2$ between high- and low-FI regimes is broadly similar across maturities—32 p.p. at 2 years, 37 at 5 years, and 21 at 10 years (average: 30 p.p.)—even though the baseline $R^2$ under high FI ranges from 84% to 33%.[35]

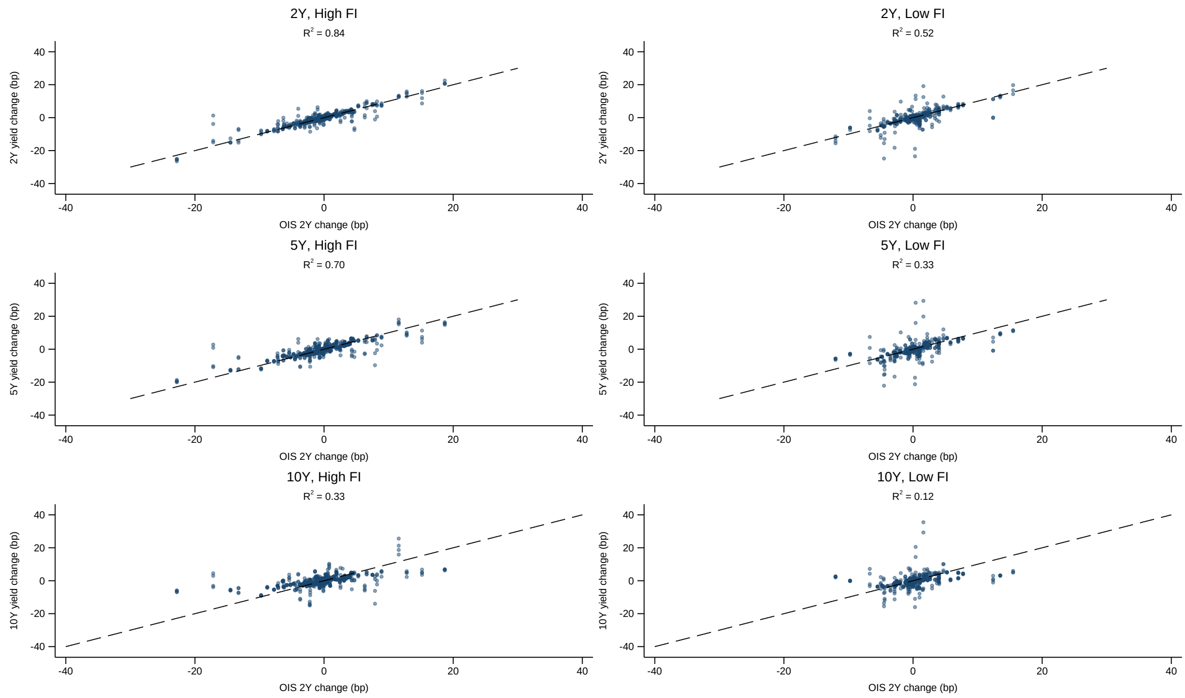
**Regression-based quantification.** For each maturity $m$, we estimate the following pass-through regression

$$\Delta y^m_{c,t} = \alpha_c + \beta^m \, \Delta \text{OIS}_{2Y,t} + \varepsilon^m_{c,t},$$

[34] Appendix A.4 replicates the pass-through analysis for the pre-crisis period.

[35] While a constant absolute $R^2$ drop does not mechanically imply a constant variance injection (because $R^2$ is a nonlinear function of the signal-to-noise ratio), this pattern is consistent with fragmentation introducing a substantial idiosyncratic component at each maturity, as the segmentation mechanism would predict.

**Figure 7:** Pass-through of ECB policy shock across the yield curve




*Notes:* Each point represents one ECB Governing Council meeting (268 meetings, January 1999 – December 2019). The horizontal axis is the change in the 2-year OIS rate (basis points) during the monetary event window from Altavilla et al. (2019); the vertical axis is the corresponding change in the sovereign yield for Germany, France, Italy, or Spain. The dashed line is the 45-degree line (perfect pass-through). $R^2$ values are from pooled OLS with heteroskedasticity-robust standard errors. High FI: meetings in quarters where the financial integration indicator exceeds the sample median; Low FI: at or below the median.

where $\Delta y^m_{c,t}$ is the change in the sovereign yield of country $c$ at maturity $m$ during ECB meeting $t$, $\Delta\text{OIS}_{2Y,t}$ is the corresponding change in the 2-year OIS rate, and $\alpha_c$ are country fixed effects. If the common policy surprise passes through uniformly, the residuals $\hat{\varepsilon}^m_{c,t}$ should be small.

We then test whether the magnitude of the idiosyncratic component depends on financial integration by estimating

$$|\hat{\varepsilon}^m_{c,t}| = \varphi_0 + \varphi_1 \, FI_t + u_{c,t},$$

where $FI_t$ is the financial integration indicator (either a binary variable equal to one when the indicator exceeds the sample median, or the continuous standardized measure), and standard errors are clustered at the meeting level to account for within-meeting correlation across countries.

Table 2 reports the results. Panel A, which uses the cross-sectional standard deviation as the dependent variable, shows the expected negative sign but the coefficients are not statistically significant at conventional levels—unsurprising given that each observation collapses four countries into a single dispersion measure, severely limiting power. Panel B, which exploits the full set of country × meeting observations, provides substantially more power. The coefficient $\varphi_1$ is negative across all specifications and statistically significant under the continuous FI measure: under the continuous FI measure, a one-standard-deviation increase in financial integration reduces the absolute pass-through residual by 4.5 basis points at the 2-year maturity

**Table 2:** Effect of financial integration on cross-country yield dispersion

| | Binary FI | | | Continuous FI | | |
|---|---|---|---|---|---|---|
| | 2-year | 5-year | 10-year | 2-year | 5-year | 10-year |
| *Panel A: Cross-sectional standard deviation of yield changes* | | | | | | |
| FI | -0.515 | -0.495 | -0.398 | -2.268 | -2.434 | -1.416 |
| | (0.313) | (0.381) | (0.349) | (1.462) | (1.833) | (1.717) |
| Meetings | 268 | 268 | 268 | 268 | 268 | 268 |
| *Panel B: Absolute residuals from pass-through regression* | | | | | | |
| FI | -0.632*** | -0.569** | -0.383 | -4.510*** | -4.259** | -3.662** |
| | (0.213) | (0.279) | (0.313) | (1.330) | (1.653) | (1.855) |
| Observations | 948 | 951 | 954 | 948 | 951 | 954 |

*Notes*: Panel A regresses the cross-sectional standard deviation of sovereign yield changes (across Germany, France, Italy, and Spain) on each meeting day on the FI indicator; standard errors are clustered at the quarterly level. Panel B regresses the absolute residuals from a pooled pass-through regression (with country fixed effects) on FI; standard errors are clustered at the meeting level. Binary FI equals one when the quarterly FI indicator exceeds the sample median. Continuous FI is the standardized FI indicator. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

($p < 0.01$), 4.3 at 5 years ($p < 0.05$), and 3.7 at 10 years ($p < 0.05$). These results are consistent with the interpretation that financial integration shapes the uniformity of the interest rate pass-through: when markets are integrated, the ECB's policy action transmits in a coordinated manner across member states; when markets are fragmented, country-specific wedges cause the common policy surprise to be a noisier proxy for the monetary impulse actually faced by each economy, attenuating the estimated average response. Higher integration reduces these wedges, strengthening the link between the common shock and national financing conditions, which our interaction specification captures.

The pass-through evidence documents one channel: integration makes the yield response to ECB actions more uniform. But uniformity of yields is neither necessary nor sufficient for amplification of real outcomes—what matters is whether the pass-through ultimately reaches credit conditions, investment, and consumption. Does the amplification of HICP and GDP operate primarily through the bond market (as the pass-through evidence would suggest), or do other financial market segments play a distinct role? To answer this question, we decompose the composite FI indicator into its four constituent sub-indices.

### 4.4.2 Channel Decomposition

The ECB also publishes four price-based sub-indices capturing integration in the money market, bond market, equity market, and banking market.[36] To assess which market segment drives the amplification, we replace the quantity-based composite FI indicator with each price-based sub-index in turn, re-estimating the baseline specification in Equation 2.

Figure 8 reports the impulse responses of HICP evaluated at the 25th and 75th percentiles of each sub-index; Figure 9 reports the corresponding GDP results. Two sub-indices stand out. *Banking market integration* shows the strongest association with GDP amplification: the

[36] The sub-indices are from the ECB's "Financial Integration and Structure in the Euro Area" statistical report (Hoffmann, Kremer and Zaharia, 2020). Each sub-index measures the degree of cross-country price convergence within the respective market segment. We collapse the monthly sub-indices to quarterly frequency to match our panel.

interaction coefficient is statistically significant from impact through horizon 7, with $t$-statistics exceeding 3.5 at the peak (horizon 4). This is consistent with a credit supply channel: when banking markets are integrated, cross-border bank lending ensures that ECB rate changes translate into uniform credit expansion across member states; under segmentation, peripheral banks face tighter balance sheet constraints and cannot expand lending as freely, weakening the output response (Peek and Rosengren, 1997; Cetorelli and Goldberg, 2012; Bruno and Shin, 2015; Kalemli-Ozcan, Papaioannou and Peydró, 2013). *Equity market integration* shows the strongest association with HICP amplification, with significant interaction coefficients from horizon 2 through 13 ($t$-statistics exceeding 2.5 at horizons 6–8). This is consistent with a portfolio rebalancing channel: when equity markets are integrated, rate changes produce uniform asset valuation gains across countries, generating wealth effects on consumption and reinforcing the demand stimulus that feeds into prices (Bernanke and Kuttner, 2005; Fratzscher, Lo Duca and Straub, 2016).

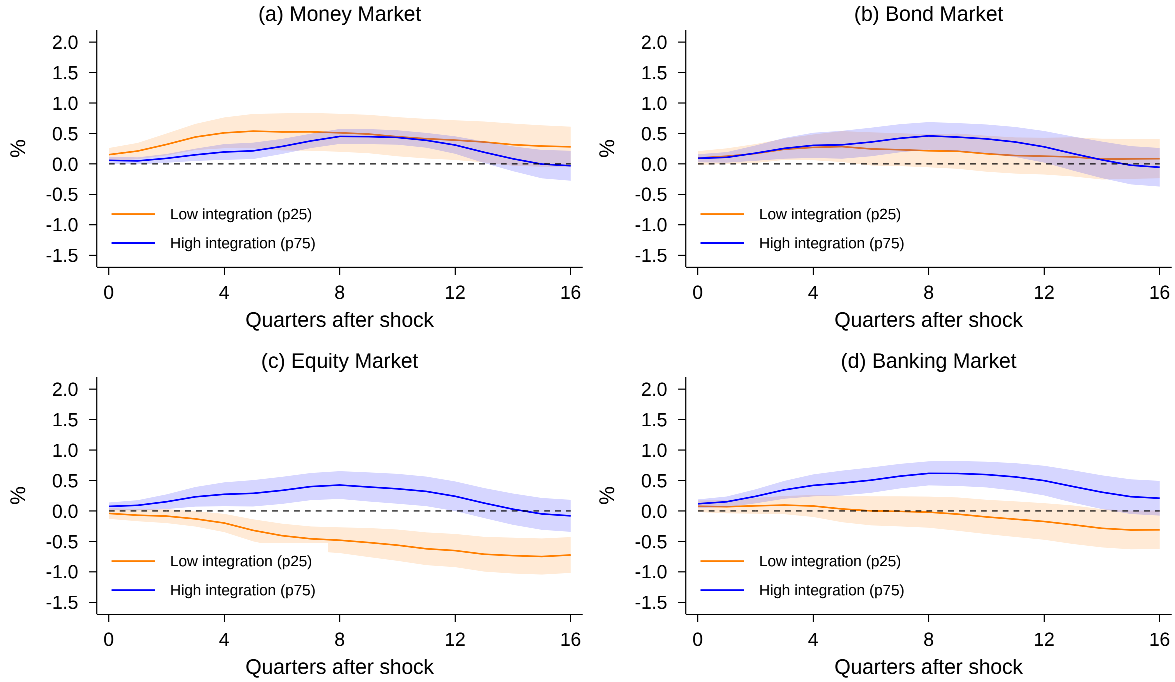
By contrast, money market integration produces no significant amplification for either variable, and bond market integration—the dimension captured by the pass-through scatter plots in Section 4.4.1—contributes moderately to GDP amplification at medium horizons but is not significant for HICP. The channel decomposition thus resolves the question raised by the pass-through analysis: while bond market integration—the channel most directly connected to the yield pass-through documented above—contributes moderately to GDP amplification, it is *not* the dominant driver for either outcome. Instead, the amplification of GDP is most strongly associated with banking integration (consistent with a credit supply channel), and the amplification of HICP is most strongly associated with equity integration (consistent with a portfolio rebalancing and wealth channel).

Importantly, no single sub-index fully subsumes the composite FI indicator. In horse-race specifications that include both the composite FI interaction and a sub-index interaction simultaneously, the composite FI coefficient is attenuated but remains significant for most horizons.[37] This suggests that the composite FI indicator captures a multi-dimensional integration phenomenon: the amplification operates through multiple financial channels—credit supply, portfolio rebalancing, and interest rate pass-through—rather than through any single market segment. Because the sub-indices are correlated, the one-at-a-time decomposition attributes variation that may be shared across market segments; the results should therefore be interpreted as revealing the strongest associations rather than independent causal contributions.

These findings are consistent with the extended model in Appendix C (Section C.6), which shows that augmenting the standard interest rate pass-through channel with direct financial channels—credit supply (banking integration) and portfolio rebalancing (equity integration)—preserves the linear interaction structure while generating a richer amplification that reflects the contribution of all three channels. The banking–GDP and equity–HICP pattern is consistent with the credit channel operating primarily through investment and the portfolio channel through consumption, which have different relative impacts on output versus prices along the Phillips curve.

[37] Appendix A.6 reports the full horse-race results.

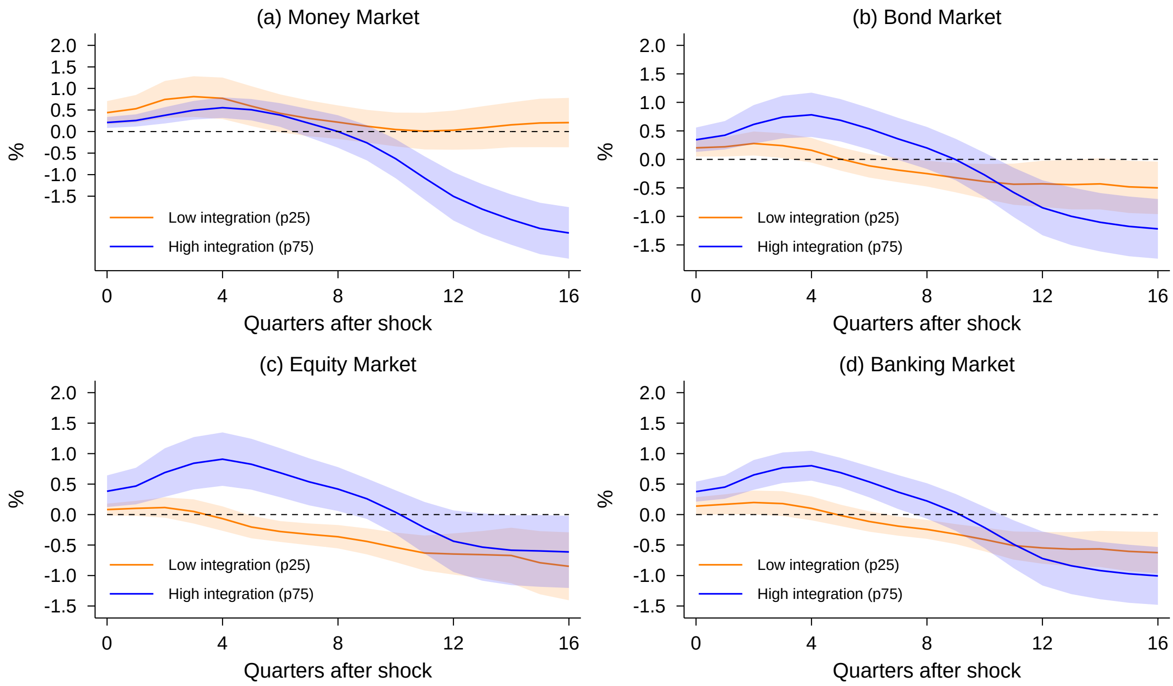
**Figure 8:** Channel decomposition: HICP responses by FI sub-index




*Notes:* Each panel replaces the composite FI indicator in Equation 2 with the corresponding ECB price-based sub-index. Impulse responses (in %) to a one-standard-deviation expansionary monetary policy shock, evaluated at the 25th percentile (orange) and 75th percentile (blue) of each sub-index. Shaded areas denote 68% confidence intervals based on Driscoll-Kraay standard errors.

The results so far paint a consistent picture: financial integration amplifies monetary transmission through multiple, complementary channels. A natural concern is that this finding may be driven by the specific empirical choices we have made—the sample period, the shock construction, the functional form of the interaction, or the co-movement of integration with cyclical conditions. We address these concerns systematically in Section 5.

**Figure 9:** Channel decomposition: GDP responses by FI sub-index




*Notes:* See notes to Figure 8. Results shown for real GDP.

## 5 Robustness and Identification

A central concern with our identification is that financial integration may co-move with the business cycle, so that the state-dependent local projections capture cyclical heterogeneity in monetary transmission rather than the structural effect of integration. As discussed in Section 2.1, the FI indicator exhibits substantial variation that is not reducible to standard business-cycle fluctuations—rising through the 2001 recession and declining most sharply during the sovereign debt crisis rather than during earlier business-cycle downturns. We nonetheless address this concern directly through multiple complementary tests: (i) augmenting the specification with an interaction between monetary shocks and an output gap indicator, (ii) estimating on a pre-2007 sample that excludes all crisis episodes entirely, (iii) down-weighting crisis-contaminated observations using the variance down-weighting method of Lenza and Primiceri (2022), and (iv) removing the crisis-period dummies from the baseline specification. The amplification effect of financial integration is robust to all four tests, indicating that our results are not driven by a small number of crisis episodes or by business-cycle comovement.

We organize the robustness analysis into three groups. The first addresses identification directly, testing whether competing state variables displace financial integration, whether the monetary shock is orthogonal to the integration state, whether the linear interaction is restrictive, and whether the results scale across alternative threshold definitions. The second tests sensitivity to data and sample choices. The third considers alternative specifications and shock measures.

### 5.1 Identification Tests

**Competing state-variable interactions.** Because financial integration varies only over time (not across countries), time fixed effects would absorb the shock–FI interaction, and any omitted time-varying confounder could bias $\delta_h$. The most prominent candidates are the business-cycle position, sovereign risk, and the output gap. We augment Equation 2 with the interaction between the monetary shock and each candidate—individually and jointly in a kitchen-sink specification—following Cloyne, Jordà and Taylor (2023). For HICP, $\delta_h$ remains negative and significant at all peak horizons across all augmented specifications. Adding the recession interaction *strengthens* $\delta_h$ (the $t$-statistic at $h = 8$ rises from 2.7 to 3.4). Adding the sovereign spread—a particularly powerful control because it is country $\times$ time varying—changes $\delta_h$ by less than 3% of its baseline value at peak horizons ($h = 8$–12), while the spread interaction itself is insignificant. In the kitchen-sink specification including all three competing interactions, $\delta_h$ remains significant at peak horizons despite wider standard errors. For GDP, $\delta_h$ is robust to the recession, spread, and output gap controls individually at the short-to-medium horizons ($h = 1$–5) where the amplification effect is economically meaningful. However, in the joint kitchen-sink specification, the GDP interaction coefficient loses significance at all reported horizons, reflecting the limited separate identification of multiple time-varying-only shock interactions. Importantly, the amplification of monetary policy transmission to consumer prices—the ECB's primary mandate variable—remains robust across all augmented specifications, including the kitchen-sink. The greater sensitivity of the GDP results is expected: the output gap interaction absorbs variation that overlaps with the FI interaction, and this overlap is more acute when the dependent variable is itself a cyclical aggregate (GDP)—whose response is directly related to the output gap—than when the dependent variable is a price-level variable (HICP) where the FI transmission channel operates primarily through sustained demand pressure. Appendix B.1.1 reports full results.

**Controlling for boom-bust states.** To address the business-cycle concern, we augment Equation 2 with an additional interaction between monetary shocks and a binary output gap indicator (using both one-sided and two-sided Hodrick–Prescott filters), replacing the continuous FI measure with a binary dummy (above/below median) to facilitate interpretation of the multiple interaction terms. This allows us to control for cyclical heterogeneity in the estimates and isolate the independent role of financial integration. The results, shown in Appendix B.1.2, support our main conclusion: monetary policy transmission is stronger in periods of higher financial integration.

**Placebo state variables.** A complementary test asks whether the amplification is specific to financial integration or could be generated by any persistent EA-aggregate time series entering the interaction. We re-estimate Equation 2 replacing FI with three placebo state variables: (i) the EA average log GDP level, (ii) the HP-filtered EA output gap, and (iii) the EA average long-term interest rate. For HICP, only financial integration produces sustained significance: the FI interaction is significant at the 90% confidence level at horizons $h = 2$–13, peaking at $|t| = 2.89$ ($h = 9$), while all three placebos remain insignificant at the peak horizons. For GDP, FI again dominates, though the EA GDP placebo also shows significance at short horizons—expected given the correlation between aggregate GDP and both FI and country-level output. We also conduct a permutation test, randomly shuffling the FI time series across quarters (200 draws) to obtain the distribution of the interaction coefficient under the null. At the HICP peak ($h = 9$), the baseline coefficient falls outside 92.5% of the permutation distribution ($p = 0.075$). Appendix B.1.3 reports the full comparison.

**Shock orthogonality across financial integration regimes.** A necessary condition for our identification is that the monetary shock is distributionally similar across FI regimes—otherwise, the interaction effect could reflect differential shock intensity rather than differential transmission. We collapse the panel to 84 quarterly observations and compare the shock distribution in high- versus low-FI quarters (median split). Four complementary tests—a two-sample $t$-test ($p = 0.643$), a variance ratio $F$-test ($p = 0.408$), a Kolmogorov–Smirnov test ($p = 0.991$), and a Wilcoxon rank-sum test ($p = 0.862$)—all fail to reject equality at any conventional significance level. The correlation between the shock and FI is $\rho = 0.008$ ($p = 0.940$). Shock magnitude ($|\widehat{\Delta i}_t|$) is also indistinguishable across regimes ($p = 0.481$). These results provide no evidence that the ECB delivers systematically different shocks during periods of high or low financial integration. Appendix B.1.4 provides full details.

**Nonlinearity of the financial integration interaction.** The linear interaction assumes that the marginal amplification effect of FI is constant. Following Tenreyro and Thwaites (2016), we test this by adding a quadratic interaction $\widehat{\Delta i}_t \times FI^2_{t-1}$ to the specification. For HICP, the quadratic term is insignificant at all 17 horizons (the largest $|t| = 1.16$), providing no evidence against the linear specification. For GDP, the quadratic term is negative and significant at 13 of 17 horizons, indicating that the marginal amplification effect of FI on the expansionary IRF increases at higher levels of integration (the coefficient polynomial $\beta_h + \delta_h FI + \delta^2_h FI^2$ is concave, which translates into a convex reported expansionary IRF). However, the implied differences in IRFs between the 75th and 25th percentiles of FI are *larger* under the quadratic specification at the relevant horizons, meaning the nonlinearity strengthens rather than weakens the amplification finding. Appendix B.1.5 reports the full set of quadratic coefficients and implied IRFs.

**Threshold robustness.** Because FI enters Equation 2 continuously, the IRF at any integration level is a direct function of the estimated coefficients, and no re-estimation is needed to evaluate regime differences at alternative thresholds. We compare four pairs—p20/p80, p25/p75 (baseline), p30/p70, and p33/p67—and find that the unsmoothed peak horizon is identical across all pairs ($h = 9$ for HICP and $h = 4$ for GDP), as expected from the proportionality of the linear specification, and that regime differences scale exactly in proportion to the FI gap between thresholds. A reader who prefers terciles rather than quartiles would see the same pattern at 58% of the baseline magnitude; one who prefers more extreme thresholds (p20/p80) would see 111%. Appendix B.1.6 provides the complete comparison.

## 5.2 Sensitivity to Data and Sample Choices

**Excluding crisis episodes.** To further ensure that our results are not an artifact of the state of the business cycle, we estimate the impulse responses on a truncated sample ending in Q2 2007, thereby excluding both the global financial crisis and the EA sovereign debt crisis. This removes the periods where financial integration and monetary transmission are most likely to be influenced by severe macro-financial stress. As reported in Appendix B.2.1, the amplification effect persists: monetary policy transmission remains significantly stronger when financial integration is higher, even outside crisis periods. Despite the smaller sample (34 quarters, versus 84 in the baseline), the amplification coefficients remain statistically significant at the 90% confidence level.

**Crisis down-weighting.** The pre-2007 sample test above removes all post-crisis data, including the recovery period during which FI partially rebounded. A complementary approach retains the full sample but down-weights crisis-contaminated observations using the variance down-weighting method of Lenza and Primiceri (2022), already applied to the COVID period in Appendix B.2.4. We define an observation as crisis-contaminated at horizon $h$ if any quarter in its dependent-variable window $[t-1, t+h]$ overlaps with the crisis period (Q3 2007–Q2 2013). This approach preserves the panel's autocorrelation structure and avoids the boundary issues inherent in sample exclusion. For each horizon, a first-pass regression estimates the variance ratio $\lambda_h$ between contaminated and normal residuals; a second-pass WLS then down-weights contaminated observations by $1/\lambda_h$. For HICP, the amplification effect is stable: the FI interaction coefficient $\delta_h$ remains significant at peak horizons, with $|t|$-statistics exceeding 2 at $h = 4$–13. For GDP, the interaction remains significant at the short-to-medium horizons ($h = 0$–7) that characterize the output amplification. These results indicate that the amplification finding is not an artifact of the crisis period. Appendix B.2.2 reports the full results and variance ratios.

**Higher frequency data.** Emerging literature suggests that time aggregation of economic activity and monetary policy shocks can affect the identification of monetary policy transmission (Buda et al., 2025). Thus, caution is needed when matching high-frequency shocks to lower-frequency macro variables like real GDP. We therefore estimate impulse responses using: monthly HICP and Industrial Production (IP) data, the same monetary shocks from Jarociński and Karadi (2020) in monthly frequency, and the price-based financial integration indicator from Hoffmann, Kremer and Zaharia (2020) available at monthly frequency. This exercise simultaneously changes the data frequency, the output measure (IP instead of GDP), and the FI indicator (price-based instead of quantity-based); Appendix B.3.4 varies the FI indicator at quarterly frequency, partially disentangling these channels. The ordering of the responses is consistent with the baseline: monetary transmission is stronger under high financial integration for both HICP and IP, although the confidence bands overlap given the imprecision inherent in the joint change of specification features.[38]

[38] Appendix B.2.3 reports the monthly-frequency results.

**Extended time frame with Lenza–Primiceri COVID handling.** The baseline sample ends in Q4 2019, excluding the COVID-19 pandemic entirely. We extend the sample to Q1 2023 and address the pandemic using the variance down-weighting approach of Lenza and Primiceri (2022). For each LP horizon $h$, we identify observations whose dependent variable window $[t-1, t+h]$ overlaps with the acute COVID period (Q1 2020–Q2 2020). A first-pass OLS regression yields residuals from which we compute the variance ratio $\lambda_h$ between contaminated and normal observations. A second-pass WLS regression then down-weights contaminated observations by $1/\lambda_h$, implemented via a manual within-transformation to preserve Driscoll–Kraay standard errors. The amplification results are closely aligned with the baseline: the peak HICP amplification is 0.50 p.p. and the peak GDP amplification is 0.47 p.p. (baseline: 0.50 p.p., ratio 0.94; see Section 4.1). The HICP peak horizon is identical to the baseline; for GDP, the peak shifts by one quarter (from $h = 4$ to $h = 5$). The first-stage $F$-statistic strengthens from 80 to 86 with the additional meetings. Appendix B.2.4 reports the full methodology and comparison with the baseline.

**Sample of countries.** We include all EA member states except the most recent member, Croatia. One may ask whether the results change if we focus only on the eleven original members that formed the EA in 1999: Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal, and Spain. In fact, the financial integration indicator is based on data for these eleven countries (except for banking and money market data). Appendix B.2.5 shows the estimated responses for the original members, indicating that the main conclusions are consistent with the full sample.

### 5.3 Alternative Specifications and Shock Measures

**Unconventional monetary policy.** Our sample spans periods of both conventional and unconventional ECB policy, including quantitative easing and long-term refinancing operations. The Jarociński and Karadi (2020) identification purges information effects but does not distinguish conventional from unconventional policy actions. To partially address this, we verify that our results hold when using the target shocks from Altavilla et al. (2019), which isolate surprises in the short-term policy rate component of ECB announcements and are therefore less influenced by unconventional policy tools.[39] We acknowledge that the post-2014 period may exhibit limited variation in the monetary shock series, and that the interaction between unconventional policy tools and financial integration warrants further investigation.

**Additional exercises.** We perform additional robustness checks by modifying the main specification in Equation 2 in several ways to ensure that the results hold. These include: (i) varying the number of lags of monetary policy shocks and of the dependent variable; (ii) omitting the time dummies that control for the crisis period; (iii) replacing the continuous measure of financial integration with a dummy based on the quantity-based composite indicator, set to one when the indicator is at or above the median and 0 otherwise; (iv) replacing

[39] Appendix B.3.5 reports the results using alternative shock series.

the quantity-based indicator with the price-based one, both from Hoffmann, Kremer and Zaharia (2020).[40]

In sum, the amplification effect of financial integration on consumer prices is robust across all specifications and identification tests we consider. As for GDP, at the short-to-medium horizons where amplification is economically meaningful, the interaction coefficient is robust to each competing control individually, to the pre-2007 subsample, to crisis down-weighting, and to all other specification checks. It attenuates only in the joint kitchen-sink specification that includes all three competing interactions simultaneously.

## 6 Conclusion

This paper shows that financial integration shapes how ECB monetary policy transmits to prices and output across the EA. A one-standard-deviation increase in financial integration amplifies the peak price response by 0.53 p.p. and the peak output response by 0.50 p.p. Comparing the 25th and 75th percentiles of integration, the difference in peak responses reaches 0.83 p.p. for HICP and 0.78 p.p. for real GDP. Under low financial integration, estimated responses of prices and output are statistically and economically insignificant; under high integration, transmission is strong and persistent.

The amplification effect is present in both core and peripheral member states, though it appears more pronounced in the periphery (0.99 versus 0.51 p.p. for HICP; 0.86 versus 0.66 p.p. for GDP)—consistent with the asymmetric nature of the fragmentation that emerged during the sovereign debt crisis.

We characterize the channels through which financial integration amplifies monetary transmission by tracing successive links in the chain. Financial integration is closely linked to the uniformity of the interest rate pass-through: the common policy surprise explains 84% of 2-year sovereign yield variation when integration is high versus 52% when it is low. Decomposing the composite indicator into its constituent market segments reveals that banking market integration shows the strongest association with output amplification, consistent with a credit supply channel, while equity market integration shows the strongest association with price amplification, consistent with a portfolio rebalancing channel. No single market segment subsumes the composite, consistent with financial integration being multi-dimensional.

Several questions remain open. Building structural theory around each of the channels we document—credit supply through banking integration, portfolio rebalancing through equity integration, and yield pass-through through bond integration—would help formalize the multi-dimensional amplification mechanism. Extending the analysis to examine how integration interacts with other sources of cross-country heterogeneity (Corsetti, Duarte and Mann, 2022) would further clarify the scope of the amplification.

Our results carry a direct policy implication: the effectiveness of the ECB's single monetary policy depends on the degree of financial integration among member states. Consistent with this view, completing the institutional architecture of financial integration in the EA—including

[40]Appendix B.3.1 (lag structure), B.3.2 (crisis dummies), B.3.3 (binary FI indicator), and B.3.4 (price-based FI indicator) report these results.

the Banking Union and the Capital Markets Union—could strengthen the transmission of monetary policy to the real economy, provided that these institutional reforms translate into the deeper cross-border market activity captured by the integration indicators used in our analysis. A complete welfare assessment would also need to weigh the financial-stability implications of deeper cross-border linkages.

## References


**Abbassi, Puriya, Falk Bräuning, Falko Fecht, and José-Luis Peydró.** 2022. "Cross-border interbank liquidity, crises, and monetary policy." *Journal of International Economics*, 139: 103657.

**Altavilla, Carlo, Luca Brugnolini, Refet S Gürkaynak, Roberto Motto, and Giuseppe Ragusa.** 2019. "Measuring euro area monetary policy." *Journal of Monetary Economics*, 108: 162–179.

**Amberg, Niklas, Thomas Jansson, Mathias Klein, and Anna Rogantini Picco.** 2022. "Five facts about the distributional income effects of monetary policy shocks." *American Economic Review: Insights*, 4(3): 289–304.

**Baele, Lieven, Annalisa Ferrando, Peter Hördahl, Elizaveta Krylova, and Cyril Monnet.** 2004. "Measuring financial integration in the euro area." ECB occasional paper.

**Bauer, Michael D., and Eric T. Swanson.** 2023. "A reassessment of monetary policy surprises and high-Frequency identification." *NBER Macroeconomics Annual*, 37(1): 87–155.

**Bayoumi, Tamim, and Barry Eichengreen.** 1992. "Shocking aspects of European monetary unification." National Bureau of Economic Research.

**Bernanke, Ben S., and Kenneth N. Kuttner.** 2005. "What Explains the Stock Market's Reaction to Federal Reserve Policy?" *The Journal of Finance*, 60(3): 1221–1257.

**Bianchi, Javier, and Louphou Coulibaly.** 2024. "Financial integration and monetary policy coordination." National Bureau of Economic Research.

**Bruno, Valentina, and Hyun Song Shin.** 2015. "Cross-border banking and global liquidity." *The Review of Economic Studies*, 82(2): 535–564.

**Buda, Gergely, Vasco M Carvalho, Giancarlo Corsetti, João B Duarte, Stephen Hansen, Afonso S Moura, Álvaro Ortiz, Tomasa Rodrigo, JV Rodríguez Mora, and G Alves da Silva.** 2025. "The short lags of monetary policy." Faculty of Economics, University of Cambridge.

**Castelnuovo, Efrem, and Giovanni Pellegrino.** 2018. "Uncertainty-dependent effects of monetary policy shocks: A new-Keynesian interpretation." *Journal of Economic Dynamics and Control*, 93: 277–296.

**Cetorelli, Nicola, and Linda S Goldberg.** 2012. "Banking globalization and monetary transmission." *The Journal of Finance*, 67(5): 1811–1843.

**Cieslak, Anna, and Andreas Schrimpf.** 2019. "Non-monetary news in central bank communication." *Journal of International Economics*, 118: 293–315.

**Cloyne, James, Òscar Jordà, and Alan M Taylor.** 2023. "State-Dependent Local Projections: Understanding Impulse Response Heterogeneity." National Bureau of Economic Research Working Paper 30971.

**Coeurdacier, Nicolas, and Hélène Rey.** 2013. "Home bias in open economy financial macroeconomics." *Journal of Economic Literature*, 51(1): 63–115.

**Corsetti, Giancarlo, Joao B Duarte, and Samuel Mann.** 2022. "One money, many markets." *Journal of the European Economic Association*, 20(1): 513–548.

**Curdia, Vasco, and Michael Woodford.** 2010. "Credit spreads and monetary policy." *Journal of Money, credit and Banking*, 42: 3–35.

**Fratzscher, Marcel, Marco Lo Duca, and Roland Straub.** 2016. "ECB unconventional monetary policy: Market impact and international spillovers." *IMF Economic Review*, 64(1): 36–74.

**Gali, Jordi, and Tommaso Monacelli.** 2008. "Optimal monetary and fiscal policy in a currency union." *Journal of international economics*, 76(1): 116–132.

**Gertler, Mark, and Peter Karadi.** 2011. "A model of unconventional monetary policy." *Journal of Monetary Economics*, 58(1): 17–34.

**Hobelsberger, Karin, Christoffer Kok Sørensen, and Francesco Paolo Mongelli.** 2022. *A tale of three crises: synergies between ECB tasks.* ECB Occasional Paper.

**Hoffmann, Peter, Manfred Kremer, and Sonia Zaharia.** 2020. "Financial integration in Europe through the lens of composite indicators." *Economics Letters*, 194: 109344.

**Hristov, Nikolay, Oliver Hülsewig, and Timo Wollmershäuser.** 2014. "The interest rate pass-through in the Euro area during the global financial crisis." *Journal of Banking & Finance*, 48: 104–119.

**Jarociński, Marek, and Peter Karadi.** 2020. "Deconstructing monetary policy surprisesthe role of information shocks." *American Economic Journal: Macroeconomics*, 12(2): 1–43.

**Jordà, Òscar.** 2005. "Estimation and Inference of Impulse Responses by Local Projections." *American Economic Review*, 95(1): 161–182.

**Jordà, Òscar, and Alan M Taylor.** 2025. "Local projections." *Journal of Economic Literature*, 63(1): 59–110.

**Kalemli-Ozcan, Sebnem, Elias Papaioannou, and José-Luis Peydró.** 2013. "Financial Regulation, Financial Globalization, and the Synchronization of Economic Activity." *The Journal of Finance*, 68(3): 1179–1228.

**Lenza, Michele, and Giorgio E Primiceri.** 2022. "How to estimate a vector autoregression after March 2020." *Journal of Applied Econometrics*, 37(4): 688–699.

**Miranda-Agrippino, Silvia, and Giovanni Ricco.** 2021. "The transmission of monetary policy shocks." *American Economic Journal: Macroeconomics*, 13(3): 74–107.

**Nakamura, Emi, and Jón Steinsson.** 2018. "High-frequency identification of monetary non-neutrality: the information effect." *The Quarterly Journal of Economics*, 133(3): 1283–1330.

**Peek, Joe, and Eric S. Rosengren.** 1997. "The International Transmission of Financial Shocks: The Case of Japan." *American Economic Review*, 87(4): 495–505.

**Tenreyro, Silvana, and Gregory Thwaites.** 2016. "Pushing on a string: US monetary policy is less powerful in recessions." *American Economic Journal: Macroeconomics*, 8(4): 43–74.

**Vari, Miklos.** 2020. "Monetary Policy Transmission with Interbank Market Fragmentation." *Journal of Money, Credit and Banking*, 52(2-3): 409–440.

**Von Borstel, Julia, Sandra Eickmeier, and Leo Krippner.** 2016. "The interest rate pass-through in the euro area during the sovereign debt crisis." *Journal of International Money and Finance*, 68: 386–402.

**Wu, Jing Cynthia, Yinxi Xie, and Ji Zhang.** 2024. "The Role of International Financial Integration in Monetary Policy Transmission." *IMF Economic Review*, 1–47.

# Appendix

## Financial Integration and the Transmission of Monetary Policy in the Euro Area

# A Additional Results

## A.1 First-Stage Regression Results

To estimate the monetary policy shocks used in Equation 2, $\widehat{\Delta i_t}$, we regress the change in the two-year German bond yield from Altavilla et al. (2019) on the high-frequency monetary policy shocks from Jarociński and Karadi (2020), using data from ECB Governing Council meeting dates only. Of the 268 meetings held between January 1999 and December 2019, three are excluded due to missing yield-change or surprise data, leaving 265 observations. The fitted values from this first-stage regression, $\widehat{\Delta i_d}$, are then summed across meeting dates within each quarter to produce the quarterly shock series used in the main analysis.

Table A.1 reports the results. The Jarociński and Karadi (2020) shock series is a strong predictor of the yield change, with an F-statistic of 79.95—well above the conventional threshold of 10—confirming instrument relevance.

**Table A.1:** First-Stage Regression

| | $\Delta i_d$ |
|---|---|
| MP surprise ($MP_d$) | 0.913*** |
| | (0.102) |
| Constant | −0.002 |
| | (0.002) |
| Observations | 265 |
| $F$-statistic | 79.95 |
| $R^2$ | 0.460 |
| Root MSE | 0.035 |

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

*Notes*: This table reports estimates of the first-stage regression $\Delta i_d = \alpha + \varphi \cdot MP_d + \varepsilon_d$, where $\Delta i_d$ is the change in the two-year German bond yield during the monetary event window from Altavilla et al. (2019), and $MP_d$ is the monetary surprise from Jarociński and Karadi (2020). The regression uses ECB Governing Council meeting dates only. Robust standard errors in parentheses.

## A.2 Local Projections Estimates

Table A.2 and Table A.3 report the raw coefficients from estimating Equation 2 for HICP and real GDP. These coefficients precede the transformations applied in the main text: (i) rescaling the shock to one standard deviation, (ii) reversing its sign to an expansionary impulse, (iii) evaluating the response at a given level of financial integration, and (iv) applying the centered moving average following Jordà and Taylor (2025). The tables therefore report responses to a 100-basis-point contractionary monetary policy shock. The "Instrumented shock" coefficient corresponds to the effect evaluated at $FI = 0$—outside the empirical support of the data—and is not interpretable as an average effect; the interaction term $\delta_h$ is invariant to the centering of FI, and all economically interpretable quantities in the main text are evaluated at observed percentiles. The "Instrumented shock $\times$ FI" term measures how a unit increase in financial integration shifts the impact of a 100-basis-point shock at each horizon.

**Table A.2:** Impulse Responses of HICP Before Adjustments Across Quarters

| Quarter | 0 | 1 | 4 | 8 | 12 | 16 |
|---|---|---|---|---|---|---|
| Instrumented shock | -0.29 | 4.15 | 12.63 | 27.03* | 28.83 | 27.42 |
| | (3.381) | (4.273) | (8.771) | (15.076) | (19.104) | (20.113) |
| Instrumented shock $\times$ FI | -4.38 | -18.53 | -56.50** | -102.34*** | -100.45** | -83.57* |
| | (8.791) | (12.648) | (24.985) | (37.573) | (47.014) | (49.850) |
| Observations | 1,501 | 1,482 | 1,425 | 1,349 | 1,273 | 1,197 |
| Within $R^2$ | 0.652 | 0.497 | 0.317 | 0.337 | 0.356 | 0.358 |
| Lag-aug controls | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Country FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

*Notes*: This table reports estimates of $y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \beta_h \cdot \widehat{\Delta i}_t + \delta_h \cdot \left(\widehat{\Delta i}_t \times FI_{t-1}\right) + \zeta_h \cdot FI_{t-1} + \sum_{k=1}^{K} \gamma_{k,h} X_{i,t-k} + \boldsymbol{\mu}_h' \mathbf{D}_t + \varepsilon_{i,t+h}$, where $y_{i,t}$ is the log of HICP for country $i$ at time $t$, and $\alpha_{i,h}$ are country fixed effects. The term $\widehat{\Delta i}_t$ denotes the series of monetary policy shocks estimated following the procedure explained in Section 2.2. $FI_{t-1}$ is the lagged measure of financial integration (i.e., the quantity-based composite indicator). The term $X_{i,t-k}$ denotes the control variables that include four lags of the outcome variable and one lag of the monetary shock. The vector $\mathbf{D}_t$ collects three contemporaneous dummies that control for the crisis episodes. Driscoll–Kraay standard errors in parentheses.

For HICP, the interaction coefficient is small and insignificant at short horizons (−4.38% at quarter 0, −18.53% at quarter 1) but grows steadily in magnitude as the horizon lengthens. By quarter 4 the coefficient reaches −56.50% (significant at the 5% level), and by quarter 8 it reaches −102.34% (significant at the 1% level). The effect remains economically large at the longest horizon: −83.57% at quarter 16, significant at the 10% level. The full horizon-by-horizon detail is in Table A.2.

**Table A.3:** Impulse Responses of real GDP Before Adjustments Across Quarters

| Quarter | 0 | 1 | 4 | 8 | 12 | 16 |
|---|---|---|---|---|---|---|
| Instrumented shock | 5.08 | 13.22** | 22.82*** | 18.57** | 3.73 | 5.57 |
| | (3.183) | (6.301) | (8.090) | (9.043) | (12.719) | (14.981) |
| Instrumented shock × FI | -27.35*** | -62.80** | -107.23*** | -68.84* | 10.43 | 11.60 |
| | (9.680) | (26.360) | (30.309) | (34.631) | (51.372) | (55.602) |
| Observations | 1,497 | 1,478 | 1,421 | 1,345 | 1,269 | 1,193 |
| Within $R^2$ | 0.181 | 0.281 | 0.418 | 0.338 | 0.317 | 0.359 |
| Lag-aug controls | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Country FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

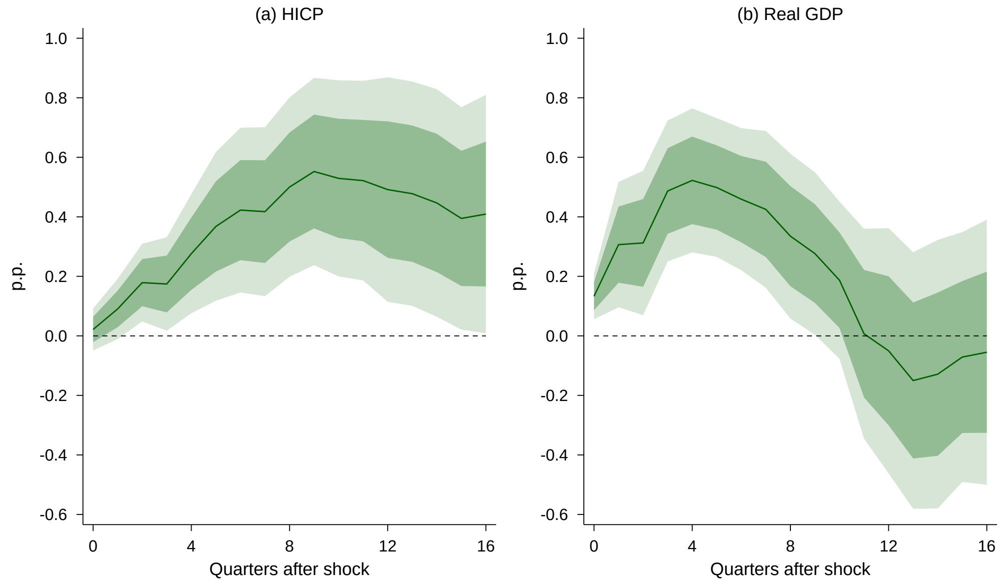
*Notes*: This table reports estimates of $y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \beta_h \cdot \widehat{\Delta i}_t + \delta_h \cdot \left(\widehat{\Delta i}_t \times FI_{t-1}\right) + \zeta_h \cdot FI_{t-1} + \sum_{k=1}^{K} \gamma_{k,h} X_{i,t-k} + \boldsymbol{\mu}_h' \mathbf{D}_t + \varepsilon_{i,t+h}$, where $y_{i,t}$ is the log of real GDP for country $i$ at time $t$, and $\alpha_{i,h}$ are country fixed effects. The term $\widehat{\Delta i}_t$ denotes the series of monetary policy shocks estimated following the procedure explained in Section 2.2. $FI_{t-1}$ is the lagged measure of financial integration (i.e., the quantity-based composite indicator). The term $X_{i,t-k}$ denotes the control variables that include four lags of the outcome variable and one lag of the monetary shock. The vector $\mathbf{D}_t$ collects three contemporaneous dummies that control for the crisis episodes. Driscoll–Kraay standard errors in parentheses.

For real GDP, the interaction coefficient is negative and significant from the outset: −27.35% at quarter 0 (1% level), rising to −62.80% at quarter 1 (5% level) and peaking at −107.23% at quarter 4 (1% level). The amplification then fades—−68.84% at quarter 8 (10% level)—and reverses sign by quarters 12 and 16, where the coefficients are positive but statistically insignificant. The pattern is consistent with an output amplification that is concentrated in the short to medium term, dissipating as the economy adjusts; the full detail is in Table A.3.

## A.3 Raw Local Projection Responses

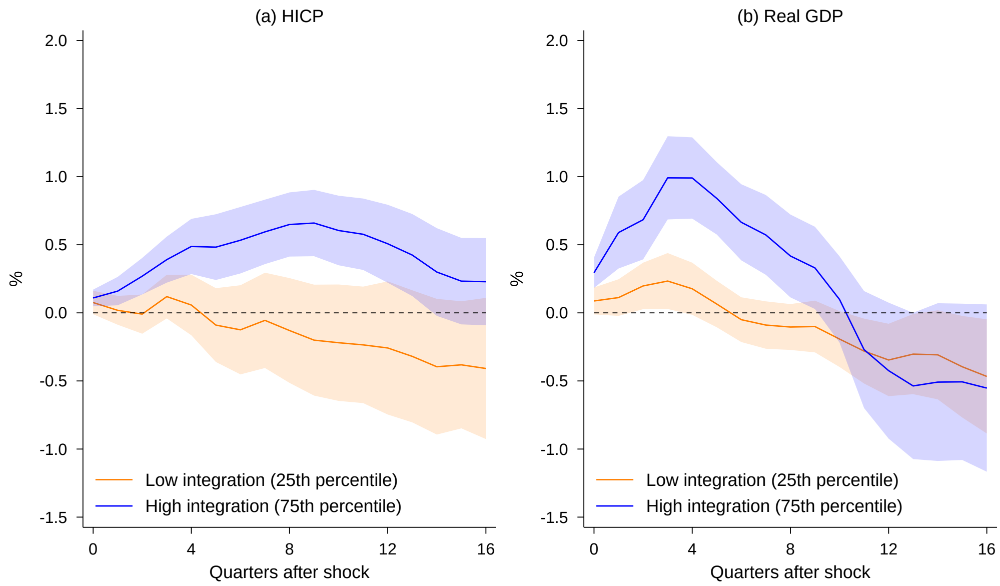
In the main analysis, as detailed in Section 3, we apply a centered moving average to the coefficient estimates from local projections, following Jordà and Taylor (2025). Because local projections estimate each horizon independently, the raw coefficients exhibit horizon-to-horizon sampling variability; this variability is naturally larger when the effective sample is smaller, as in the pre-2007 subsample. Here, we present the unsmoothed responses for transparency.

**Figure A.1:** Amplification effects of increased financial integration on HICP and real GDP responses



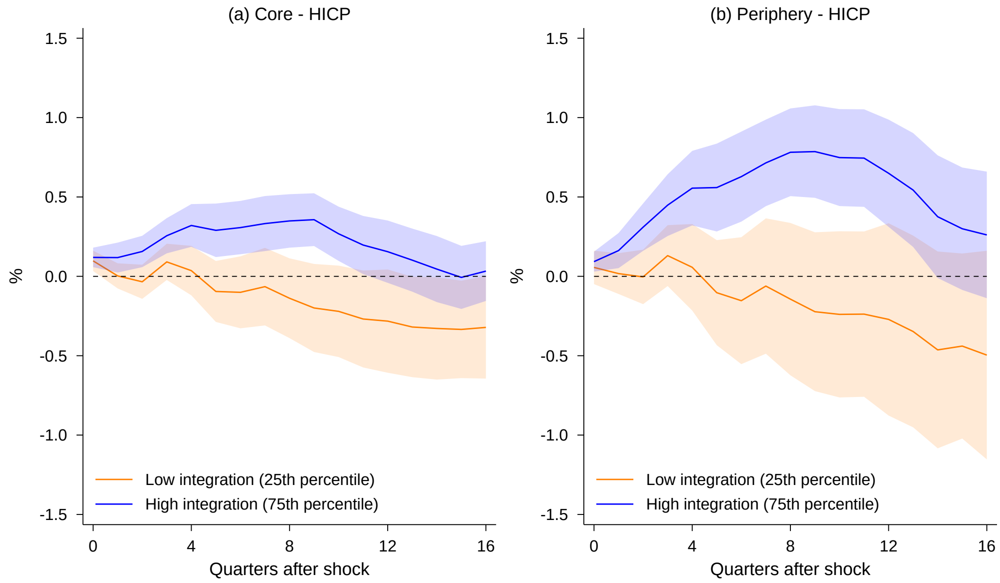

*Notes:* Amplification effects (p.p.) of a one-standard-deviation increase in financial integration on the impulse responses of HICP and real GDP. Shaded areas represent 68% and 90% confidence intervals, based on Driscoll-Kraay standard errors. These responses are obtained before the step of applying the centered moving average over the coefficients following Jordà and Taylor (2025).

**Figure A.2:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration



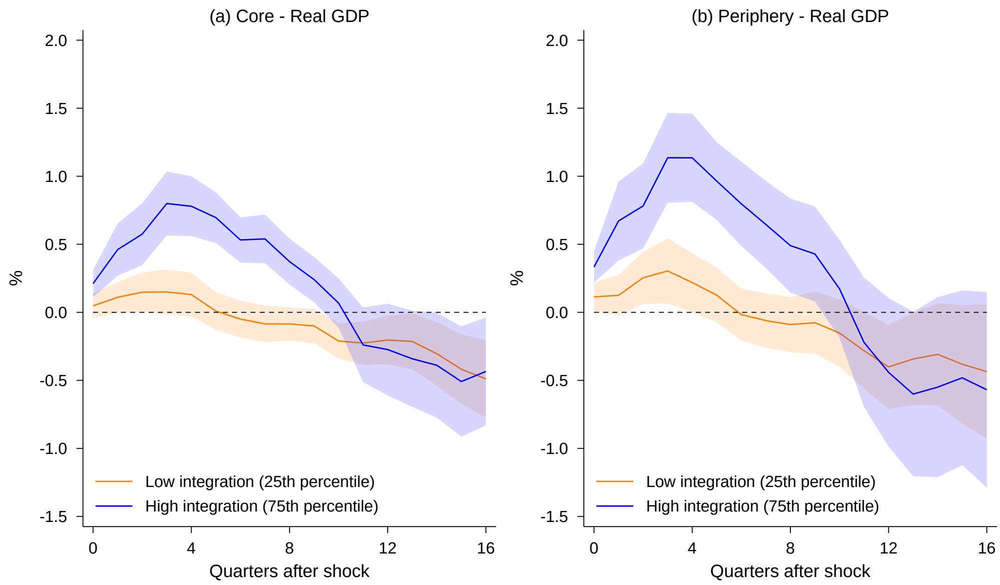

*Notes:* Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary policy shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors. These responses are obtained before the step of applying the centered moving average over the coefficients following Jordà and Taylor (2025).

**Figure A.3:** Responses of HICP to a monetary policy shock under different levels of financial integration for core and periphery




*Notes:* Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary policy shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors. These responses are obtained before the step of applying the centered moving average over the coefficients following Jordà and Taylor (2025).

**Figure A.4:** Responses of real GDP to a monetary policy shock under different levels of financial integration for core and periphery

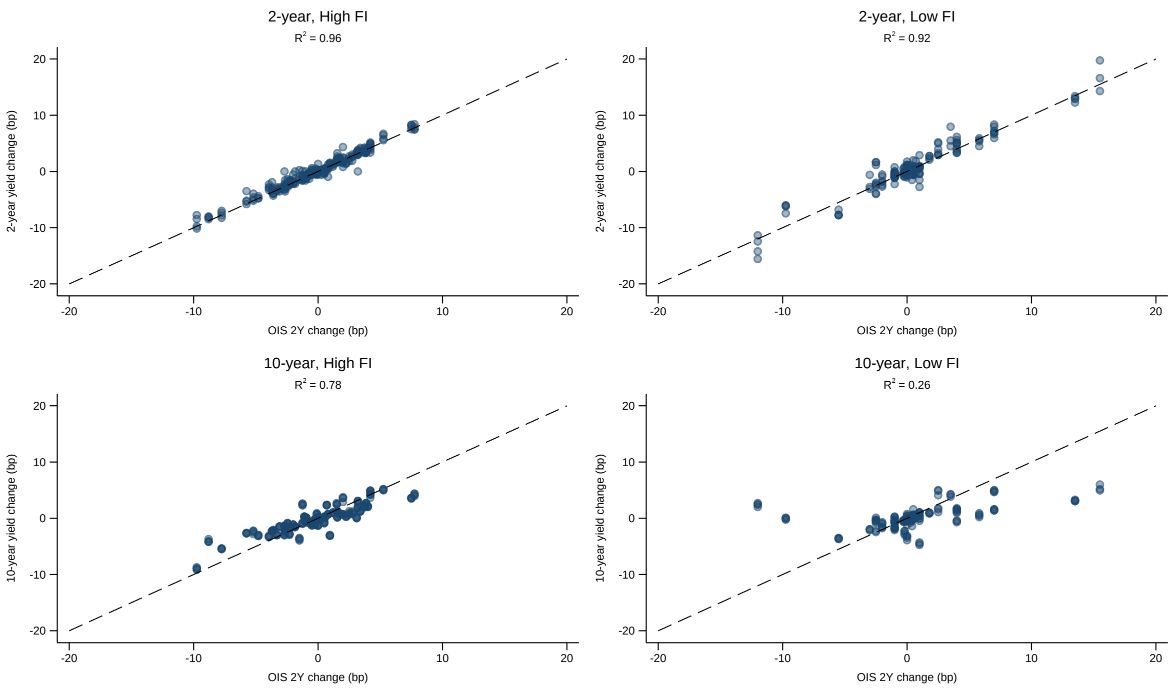



*Notes:* Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary policy shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors. These responses are obtained before the step of applying the centered moving average over the coefficients following Jordà and Taylor (2025).

## A.4 Interest Rate Pass-Through: Pre-Crisis Robustness

Figure A.5 replicates the pass-through analysis from Section 4.4.1 restricting the sample to the pre-crisis period (January 1999 to June 2007, 137 meetings). This subsample is designed so that the variation in financial integration predominantly reflects the gradual process of market integration rather than sovereign risk dynamics associated with the subsequent debt crisis. The high and low FI regimes are defined using the within-subsample median. At the 2-year maturity, the pooled $R^2$ is 0.96 under high financial integration and 0.92 under low integration—a small difference, consistent with near-complete pass-through at short maturities in both regimes. The contrast widens at longer maturities: at the 10-year maturity, the pooled $R^2$ drops from 0.78 under high FI to 0.26 under low FI, and the pass-through coefficient falls substantially. This pattern is consistent with the main finding being present before the onset of the sovereign debt crisis, though the monotonic pre-crisis trend in FI means that other trending variables cannot be fully ruled out as confounders.

**Figure A.5:** Pass-through of ECB policy shock to sovereign yields, pre-crisis period




*Notes:* Each point represents one ECB Governing Council meeting during the pre-crisis period (137 meetings, January 1999 – June 2007). The horizontal axis is the change in the 2-year OIS rate (basis points) during the monetary event window from Altavilla et al. (2019); the vertical axis is the corresponding change in the sovereign yield for Germany, France, Italy, or Spain. The dashed line is the 45-degree line (perfect pass-through). $R^2$ values are from pooled OLS with heteroskedasticity-robust standard errors. High FI: meetings in quarters where the financial integration indicator exceeds the within-subsample median; Low FI: at or below the median.

## A.5 Pass-Through Interaction with Financial Integration

As discussed in Section 4.4.1, the pass-through *coefficient* (the slope of sovereign yield changes on the OIS rate change) remains broadly stable across FI regimes; what changes is the *dispersion* around it—that is, FI shapes the uniformity of pass-through rather than its average level. Table A.4 verifies this formally: regressing sovereign yield changes on the OIS rate change interacted with FI produces insignificant interaction terms at all maturities, under both binary and continuous FI specifications. The OIS coefficient itself is significant at the 1% level in all specifications, consistent with the ECB policy surprise remaining a relevant common factor regardless of the FI regime. Standard errors are clustered at the meeting level to account for within-meeting correlation across countries.

**Table A.4:** Pass-through interaction with financial integration

| | Binary FI | | | Continuous FI | | |
|---|---|---|---|---|---|---|
| | 2-year | 5-year | 10-year | 2-year | 5-year | 10-year |
| $\Delta\mathrm{OIS}_{2Y}$ | 0.953*** | 0.757*** | 0.396*** | 0.913*** | 0.739*** | 0.422*** |
| | (0.076) | (0.090) | (0.095) | (0.055) | (0.063) | (0.073) |
| $\Delta\mathrm{OIS}_{2Y} \times$ FI | -0.031 | -0.004 | 0.034 | 0.038 | 0.034 | -0.004 |
| | (0.096) | (0.110) | (0.123) | (0.053) | (0.045) | (0.053) |
| Country FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 948 | 951 | 954 | 948 | 951 | 954 |
| $R^2$ | 0.710 | 0.527 | 0.229 | 0.711 | 0.528 | 0.229 |

*Notes*: OLS regression of sovereign yield changes on the 2-year OIS rate change and its interaction with FI, with country fixed effects. Standard errors are clustered at the meeting level. Binary FI equals one when the quarterly FI indicator exceeds the sample median. Continuous FI is the standardized FI indicator. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

### A.6 Channel Decomposition: Horse-Race Specifications

To assess whether any individual market segment subsumes the composite FI indicator, we estimate horse-race specifications that include both the composite FI interaction and a sub-index interaction simultaneously. Specifically, for each sub-index $j \in \{\text{money, bond, equity, banking}\}$, we estimate:

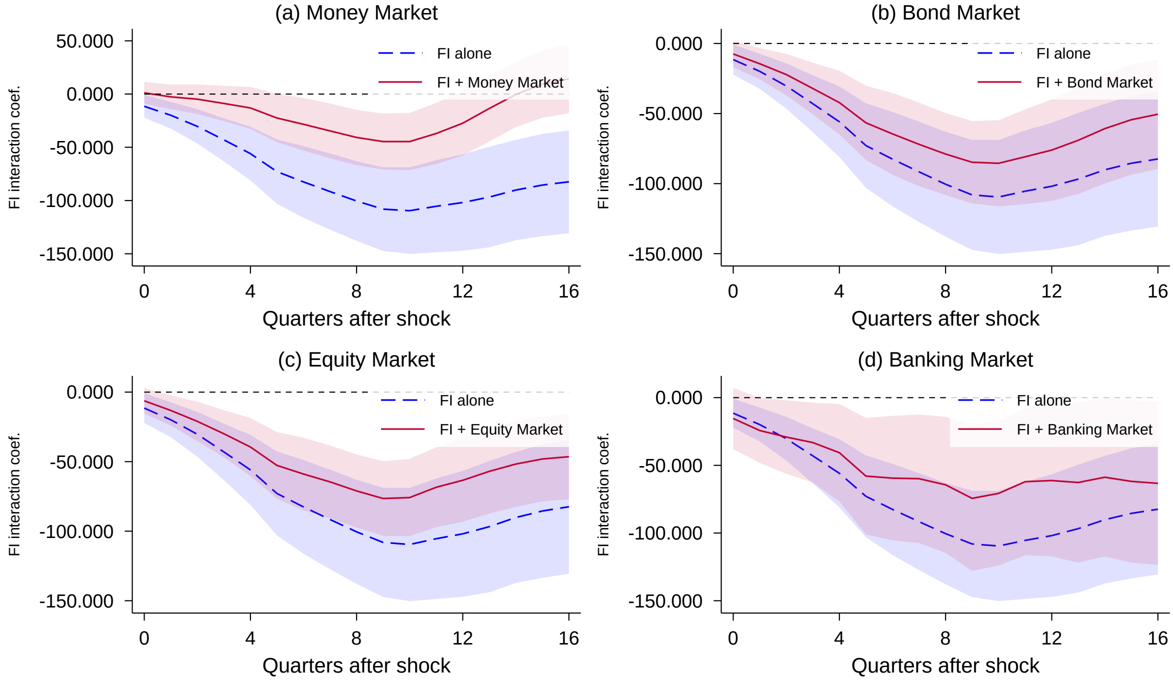
$$y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \left(\beta_h + \delta_h\, FI_{t-1} + \psi_{j,h}\, Sub^j_{t-1}\right) \widehat{\Delta i_t} + \Gamma_h(L)\mathbf{X}_{i,t-1} + \varepsilon_{i,t+h}$$

where $FI_{t-1}$ is the composite quantity-based FI indicator and $Sub^j_{t-1}$ is the price-based sub-index for market segment $j$, both lagged one quarter. The specification also includes the level effects of $FI_{t-1}$ and $Sub^j_{t-1}$ as well as three crisis-period dummies, suppressed here for brevity.

Figures A.6 and A.7 report the path of the composite FI interaction coefficient $\delta_h$ across horizons in the baseline specification (dashed blue line) and each horse-race specification that adds the sub-index interaction (solid red line). If the sub-index fully subsumes the composite, $\delta_h$ should become insignificant.

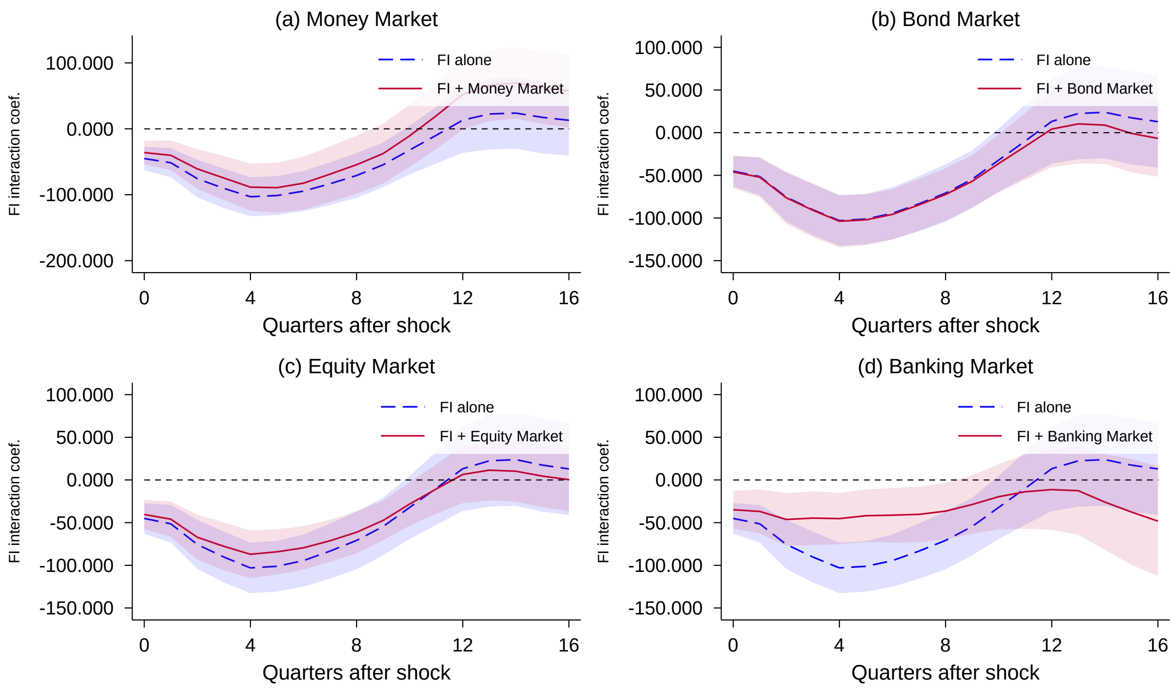
Across all four horse races, the composite FI coefficient is attenuated but remains significant at most horizons, including the peak horizons for both HICP and GDP. No single market segment fully accounts for the amplification captured by the composite indicator, consistent with financial integration operating through multiple market channels simultaneously.

**Figure A.6:** Horse-race specifications: composite FI coefficient path (HICP)




*Notes:* Each panel adds the indicated sub-index interaction to the baseline specification alongside the composite FI interaction. The dashed blue line shows the composite FI interaction coefficient $\hat{\delta}_h$ in the baseline specification (without the sub-index); the solid red line shows $\hat{\delta}_h$ when the sub-index interaction is included. Shaded areas denote 68% confidence intervals based on Driscoll-Kraay standard errors.

**Figure A.7:** Horse-race specifications: composite FI coefficient path (GDP)




*Notes:* See notes to Figure A.6. Results shown for real GDP.

# B Robustness Checks

## B.1 Identification Tests

### B.1.1 Competing state-variable interactions

A central concern with the interaction specification in Equation 2 is that financial integration varies only over time, precluding the inclusion of time fixed effects. Any omitted time-varying state variable that independently modulates monetary policy transmission could therefore bias the interaction coefficient $\delta_h$. The most prominent candidates are: (i) the business-cycle position, since FI may co-move with recessions or expansions; (ii) sovereign risk, since FI declined precisely when sovereign spreads widened during the debt crisis; and (iii) the output gap, which captures the continuous cyclical position of the economy.

We address this concern by augmenting Equation 2 with competing interaction terms. For each candidate state variable $Z$, the augmented specification adds both the interaction $\widehat{\Delta i_t} \times Z_{t-1}$ and the level $Z_{t-1}$:

$$y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \beta_h \widehat{\Delta i_t} + \delta_h \left( \widehat{\Delta i_t} \times FI_{t-1} \right) + \gamma_h \left( \widehat{\Delta i_t} \times Z_{t-1} \right) + \zeta_h FI_{t-1} + \phi_h Z_{t-1} + \Gamma_h' X_{i,t} + \varepsilon_{i,t+h} \tag{B.1}$$

If the FI interaction is absorbing the effect of $Z$, then adding $Z$ should attenuate or eliminate $\delta_h$ while $\gamma_h$ captures the omitted channel. We consider three state variables individually and jointly in a "kitchen-sink" specification: (a) a recession indicator equal to one when the EA output gap is negative (binary, time-varying only); (b) the sovereign spread, defined as country $i$'s 10-year government bond yield minus Germany's (country $\times$ time varying); and (c) the continuous output gap extracted from a Hodrick–Prescott filter (time-varying only). This approach follows Cloyne, Jordà and Taylor (2023).

**HICP results.** Table B.1 reports the FI interaction coefficient $\delta_h$ across the baseline and four augmented specifications at key horizons. Adding the recession interaction (Spec 2) *strengthens* $\delta_h$: the $t$-statistic at $h = 8$ rises from 2.7 to 3.4, indicating that once cyclical heterogeneity is purged, the structural FI effect is more precisely estimated. Adding the sovereign spread (Spec 3) barely changes $\delta_h$—within 3% of the baseline at peak horizons ($h = 8$–12)—despite introducing country $\times$ time variation that FI cannot capture and despite reducing the sample from 1,501 to 1,168 observations due to missing long-term rate data.[41] The spread interaction $\gamma_h$ is small and insignificant throughout (Table B.3). Adding the continuous output gap (Spec 4) attenuates $\delta_h$ by 25–35% at some horizons, which is expected given the overlap between two time-varying-only shock interactions that compete for the same identifying variation. In the kitchen-sink specification (Spec 5), $\delta_h$ remains negative and statistically significant at the peak horizons ($h = 5$–9, $|t| > 2$).

[41] The sample reduction from 1,501 to 1,168 reflects missing long-term sovereign yield data for some country-quarter observations. The stability of $\delta_h$ despite this 22% sample reduction suggests that neither the additional control nor the sample composition drives the result, though a formal like-for-like comparison on the restricted sample would further isolate the role of the spread control.

**Table B.1:** FI interaction coefficient $\delta_h$ across specifications — HICP

| $h$ | Baseline | + Recession | + Spread | + Gap | All |
|---|---|---|---|---|---|
| 0 | $-4.4$ (8.8) | $-7.1$ (8.7) | $-4.0$ (7.9) | $-4.5$ (9.8) | $-14.7$ (15.1) |
| 1 | $-18.6$ (12.6) | $-21.1$ (14.0) | $-17.6$ (12.2) | $-19.9$ (13.8) | $-35.1$ (25.1) |
| 4 | $-56.6$** (24.9) | $-77.2$*** (23.8) | $-53.9$** (23.7) | $-36.9$ (25.5) | $-64.7$* (38.0) |
| 8 | $-102.5$*** (37.5) | $-123.4$*** (35.8) | $-100.7$*** (35.2) | $-76.7$** (34.9) | $-103.0$** (49.7) |
| 9 | $-113.2$*** (39.2) | $-129.9$*** (39.7) | $-112.0$*** (36.5) | $-94.8$** (38.0) | $-128.2$** (54.8) |
| 12 | $-100.8$** (47.0) | $-119.8$*** (43.6) | $-98.0$** (42.6) | $-68.5$ (46.1) | $-114.1$** (56.6) |
| 16 | $-83.9$* (49.9) | $-110.0$** (44.6) | $-81.0$* (44.7) | $-47.0$ (50.5) | $-115.3$** (53.9) |
| $N$ | 1,501 | 1,501 | 1,168 | 1,501 | 1,168 |

*Notes:* Driscoll–Kraay standard errors with 4 lags in parentheses. *$p < 0.10$, **$p < 0.05$, ***$p < 0.01$. "+ Recession" adds an interaction between the shock and a recession indicator; "+ Spread" adds an interaction with the sovereign spread (country $\times$ time varying); "+ Gap" adds an interaction with the continuous output gap; "All" includes all three competing interactions simultaneously.

**GDP results.** Table B.2 reports analogous results for real GDP. At the short-to-medium horizons ($h = 0$–5) where the GDP amplification effect is economically meaningful, $\delta_h$ remains negative and significant across the baseline, recession, and spread specifications. The sovereign spread specification changes $\delta_h$ by less than 10% at peak horizons while the spread interaction is itself insignificant (Table B.4). The output gap attenuates $\delta_h$ more substantially for GDP than for HICP: at the GDP peak horizon ($h = 4$), $\delta_h$ moves from $-107.1$ to $-39.1$ when the gap interaction is added, suggesting that the output gap interaction absorbs a substantial share of the FI interaction's explanatory power for this cyclical outcome. At long horizons ($h > 8$), the GDP results are inherently less stable across all specifications, consistent with the baseline finding that FI amplifies output only temporarily. Importantly, the amplification of monetary policy transmission to consumer prices—the ECB's primary mandate variable—remains robust across all augmented specifications, including the kitchen-sink. This asymmetry between HICP and GDP robustness is expected: the output gap interaction absorbs variation that overlaps with the FI interaction, and this overlap is more acute when the dependent variable is itself a cyclical aggregate (GDP)—whose response is directly related to the output gap—than when the dependent variable is a price-level variable (HICP) where the FI transmission channel operates primarily through sustained demand pressure.

**Table B.2:** FI interaction coefficient $\delta_h$ across specifications — real GDP

| $h$ | Baseline | + Recession | + Spread | + Gap | All |
|---|---|---|---|---|---|
| 0 | −27.3*** (9.6) | −27.6*** (9.8) | −20.1** (9.4) | −14.3 (10.1) | −6.0 (16.0) |
| 1 | −62.8** (26.3) | −68.5*** (25.3) | −54.8** (24.5) | −26.0** (13.0) | −27.7 (21.2) |
| 3 | −99.8*** (29.5) | −104.8*** (28.0) | −88.4*** (25.4) | −35.8** (17.1) | −16.6 (20.2) |
| 4 | −107.1*** (30.1) | −110.2*** (29.3) | −97.9*** (27.3) | −39.1** (16.8) | −21.1 (28.0) |
| 5 | −102.2*** (29.1) | −100.5*** (29.0) | −95.6*** (26.8) | −40.4** (16.3) | −27.5 (29.1) |
| 8 | −68.7** (34.5) | −48.7 (46.0) | −58.4 (36.3) | −12.1 (28.2) | 26.0 (62.8) |
| 12 | 10.3 (51.4) | 3.0 (41.9) | 28.7 (53.9) | 103.4*** (35.7) | 98.7* (53.2) |
| $N$ | 1,497 | 1,497 | 1,168 | 1,497 | 1,168 |

*Notes:* See notes to Table B.1. Dependent variable is log real GDP.

Tables B.3 and B.4 report the competing interaction coefficients $\gamma_h$. For the sovereign spread, $\gamma_h$ is small and statistically insignificant at nearly all horizons for both outcomes, confirming that sovereign risk does not independently modulate monetary transmission once FI is controlled for. The recession interaction is positive and significant at some horizons for HICP, indicating that recessions independently dampen price transmission—but this does not absorb the FI effect. The output gap interaction is generally insignificant.

**Table B.3:** Competing interaction coefficients $\gamma_h$ at key horizons — HICP

| $h$ | Recession (Sp2) | Recession (Sp5) | Spread (Sp3) | Spread (Sp5) | Gap (Sp4) | Gap (Sp5) |
|---|---|---|---|---|---|---|
| 0 | 2.0 (2.1) | 4.2 (3.8) | 0.3 (0.3) | 0.5 (0.3) | 0.2 (0.5) | 1.0 (0.9) |
| 1 | 1.2 (3.0) | 5.9 (6.4) | 0.5 (0.5) | 0.9 (0.4) | 0.6 (1.0) | 2.1 (1.7) |
| 4 | 14.0*** (5.3) | 12.2 (10.4) | −0.1 (1.0) | 0.1 (0.9) | −2.4 (2.2) | 0.5 (3.1) |
| 8 | 13.6* (7.1) | 10.4 (13.9) | 0.3 (1.3) | 0.6 (1.4) | −2.5 (2.3) | 0.9 (3.4) |
| 12 | 15.7* (8.1) | 18.4 (14.4) | 0.7 (1.5) | 1.5 (1.7) | −1.3 (2.2) | 3.6 (3.1) |
| 16 | 21.9** (10.0) | 28.1* (15.3) | 0.5 (1.5) | 1.5 (1.9) | −2.2 (2.5) | 4.4 (2.9) |

*Notes:* Driscoll–Kraay standard errors in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. "Sp2" through "Sp5" refer to the specification number from Tables B.1–B.2.

**Table B.4:** Competing interaction coefficients $\gamma_h$ at key horizons — real GDP

| $h$ | Recession (Sp2) | Recession (Sp5) | Spread (Sp3) | Spread (Sp5) | Gap (Sp4) | Gap (Sp5) |
|---|---|---|---|---|---|---|
| 0 | 2.4 (3.0) | 0.4 (3.9) | −0.4 (0.4) | −0.5 (0.4) | 0.1 (0.6) | −0.0 (0.9) |
| 1 | 9.5 (6.3) | 5.2 (6.3) | 0.0 (0.9) | −0.3 (0.9) | −0.8 (1.3) | −0.1 (1.6) |
| 4 | 11.9 (7.9) | −3.0 (5.3) | 0.9 (1.8) | −0.3 (1.9) | −2.1** (1.3) | −2.9* (1.6) |
| 8 | −2.4 (12.8) | −9.0 (16.0) | −0.4 (2.5) | −1.5 (3.1) | 4.0 (2.5) | 1.6 (3.7) |
| 12 | 22.2* (11.6) | 13.3 (14.9) | −3.0 (2.4) | −4.2 (3.1) | 1.4 (2.7) | 2.0 (4.2) |
| 16 | 33.5*** (12.8) | 18.1 (16.6) | −3.8 (2.5) | −5.6** (3.0) | −2.6 (3.0) | −0.8 (4.1) |

*Notes:* See notes to Table B.3. Dependent variable is log real GDP.

#### B.1.2 Controlling for boom-bust states

A potential concern with our identification strategy is that the observed heterogeneity in monetary policy transmission with different levels of financial integration could partly reflect the business cycle. Our baseline sample (Q1 1999 – Q4 2019) includes episodes of severe macroeconomic stress, such as the global financial crisis and the EA sovereign debt crisis, when both financial integration and the responsiveness to monetary shocks may have co-moved with cyclical fluctuations. Without accounting for this, the estimated amplification effect could overstate the structural role of integration.

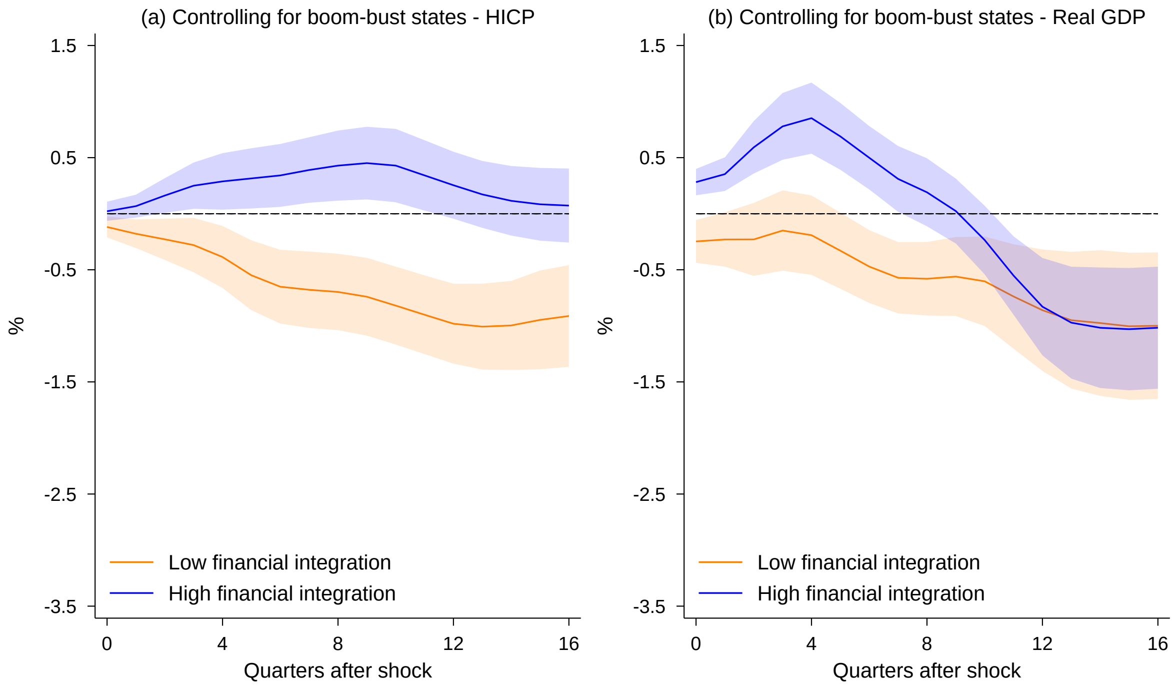
To address this concern, we extend Equation 2 by including another interaction term. This additional interaction is between the monetary policy shock series and an indicator of recession. We consider two alternative measures of economic busts for the indicator of recession:

- **Output gap dummy (one-sided HP filter)**: a dummy equal to one whenever the output gap is negative, with potential real GDP estimated using a one-sided Hodrick–Prescott (HP) filter.
- **Output gap dummy (two-sided HP filter)**: a dummy equal to one whenever the output gap is negative, with potential real GDP estimated using a two-sided HP filter.

For this robustness exercise, we depart from the baseline approach of defining high and low financial integration as the 25th and 75th percentiles. Instead, we create a binary financial integration dummy set to one when the integration level exceeds the sample median. For completeness, Appendix B.3.3 shows an exercise where the only modification to the original Equation 2 is replacing the continuous financial integration measure with a binary indicator.

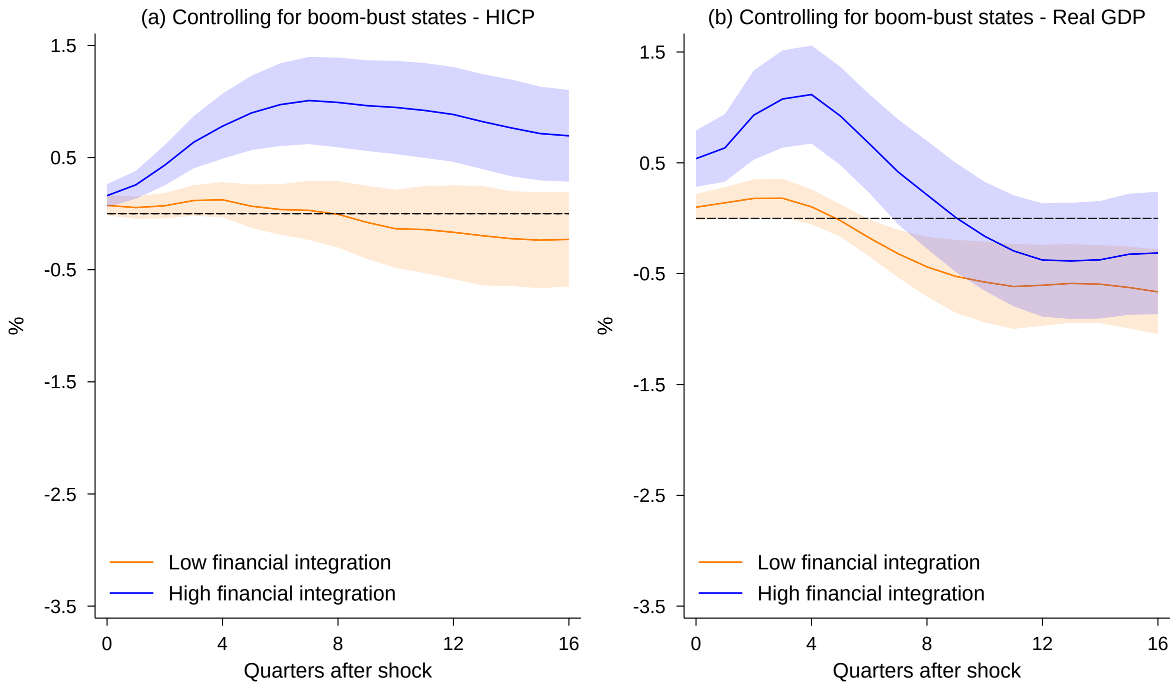
The amplification effect of financial integration survives both recession indicators, indicating that the baseline findings are not primarily driven by cyclical fluctuations captured by standard recession measures.

**Figure B.1:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Additional interaction term using output gap dummy (one-sided HP filter). Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (below median, orange line) and high (above median, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

**Figure B.2:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Additional interaction term using output gap dummy (two-sided HP filter). Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (below median, orange line) and high (above median, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

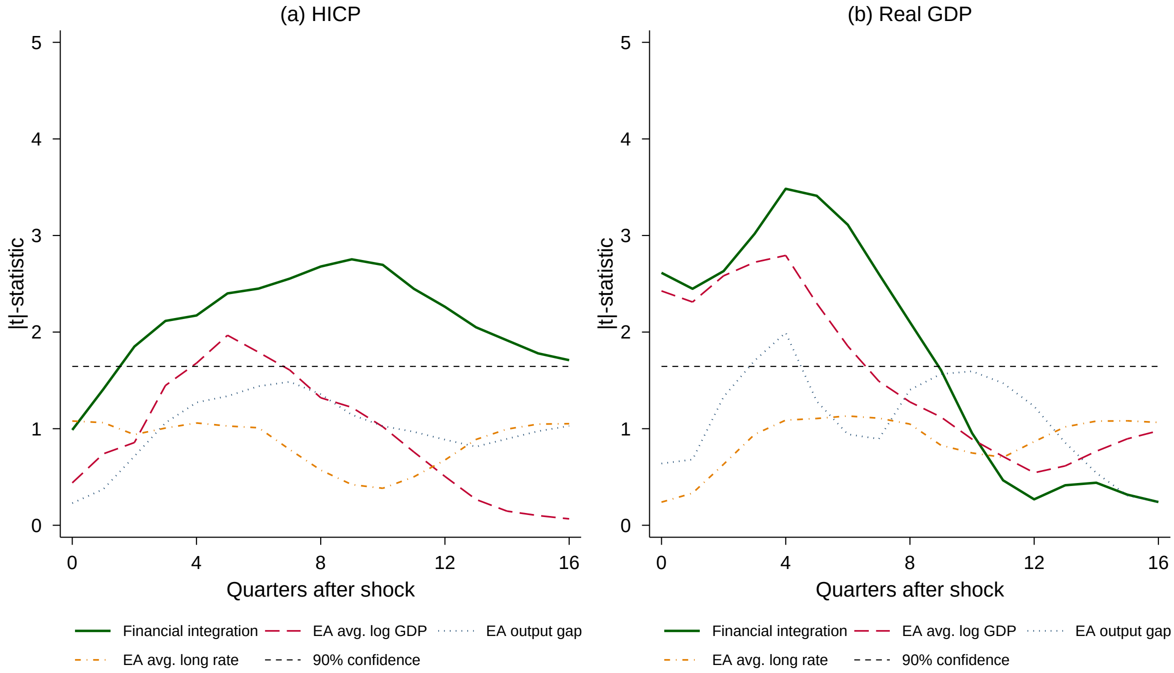
### B.1.3 Placebo state variables

Because the FI indicator varies only over time, one may worry that any sufficiently persistent EA-aggregate time series could generate a significant interaction with the monetary shock. We test this by replacing the FI measure in Equation 2 with three placebo state variables: (i) the EA average log GDP level, (ii) the HP-filtered EA output gap (smoothing parameter 1600), and (iii) the EA average long-term interest rate. All three are EA-aggregate, time-varying-only variables with substantial persistence—precisely the properties that could in principle mimic FI in the interaction.

Figure B.3 plots the absolute $t$-statistics of the interaction coefficient $\delta_h$ across horizons for the baseline FI specification and each placebo. For HICP (panel a), only the FI interaction sustains significance above the 90% confidence threshold, doing so from $h = 2$ through $h = 13$ with a peak of $|t| = 2.89$ at $h = 9$. The three placebos remain below the significance threshold at all peak horizons. For GDP (panel b), FI produces the largest $t$-statistics (peaking at $|t| = 3.55$ at $h = 4$), though the EA GDP placebo also crosses the threshold at short horizons ($h = 0$–5). This is expected: the EA GDP level is correlated with both FI and country-level output, creating a mechanical overlap when GDP is the dependent variable. The output gap and EA long-term rate placebos are insignificant at most horizons.

As a complementary test, we randomly shuffle the FI time series across the 84 sample quarters (200 draws, seed 20260224) and re-estimate the specification using OLS with country fixed effects to obtain the permutation distribution of the interaction coefficient under the null. At the HICP peak ($h = 9$), 15 of 200 permuted coefficients exceed the baseline in absolute value, yielding a permutation $p$-value of 0.075—rejecting at the 10% level but not at 5%. Because the permutation destroys all temporal dependence in FI, not only the variation relevant for identification, the effective null hypothesis is conservative.

**Figure B.3:** Placebo state variables: $|t|$-statistics of the interaction coefficient




*Notes:* Each line plots the absolute $t$-statistic of the interaction coefficient $\hat{\delta}_h$ at each horizon $h$, estimated from Equation 2 with the indicated state variable replacing financial integration. The dashed horizontal line marks the 90% confidence threshold ($|t| = 1.645$). Smoothed with a centered moving average of order 3. Driscoll-Kraay standard errors with 4 lags.

### B.1.4 Shock distribution across financial integration regimes

If the ECB delivers systematically different monetary policy shocks during periods of high or low financial integration, the interaction coefficient $\delta_h$ could reflect differential shock *intensity* rather than differential *transmission*. A necessary condition for our identification strategy is that the shock is distributionally similar across FI regimes—that is, "as good as random" with respect to the level of financial integration.

We test this by collapsing the panel to 84 unique quarters (Q1 1999–Q4 2019)—the level at which both $\widehat{\Delta i_t}$ and $FI_t$ are defined—and splitting the sample at the median FI level ($FI = 0.2988$), yielding 42 quarters per regime. Table B.5 reports descriptive statistics. The mean shock is $-0.0004$ in low-FI quarters and 0.0054 in high-FI quarters—a difference of 0.006, economically negligible relative to the shock's standard deviation of 0.057. The standard deviations are also similar (0.053 vs. 0.060).

**Table B.5:** Monetary policy shock: descriptive statistics by FI regime

| | $N$ | Mean | SD | Min | Max |
|---|---|---|---|---|---|
| Low FI (below median) | 42 | $-0.00037$ | 0.05302 | $-0.21225$ | 0.12093 |
| High FI (above median) | 42 | 0.00540 | 0.06040 | $-0.12722$ | 0.22582 |
| All quarters | 84 | 0.00252 | 0.05656 | $-0.21225$ | 0.22582 |

Table B.6 reports four formal statistical tests. The two-sample $t$-test fails to reject equal means ($p = 0.643$). The variance ratio $F$-test fails to reject equal variances ($p = 0.408$). The Kolmogorov–Smirnov test, which is sensitive to differences in the full distributional shape, yields $D = 0.095$ ($p = 0.991$), indicating that the two empirical CDFs are nearly indistinguishable. The Wilcoxon rank-sum test, which is robust to the heavy tails present in the shock series (kurtosis = 7.3), also fails to reject ($p = 0.862$). All four $p$-values exceed 0.40.

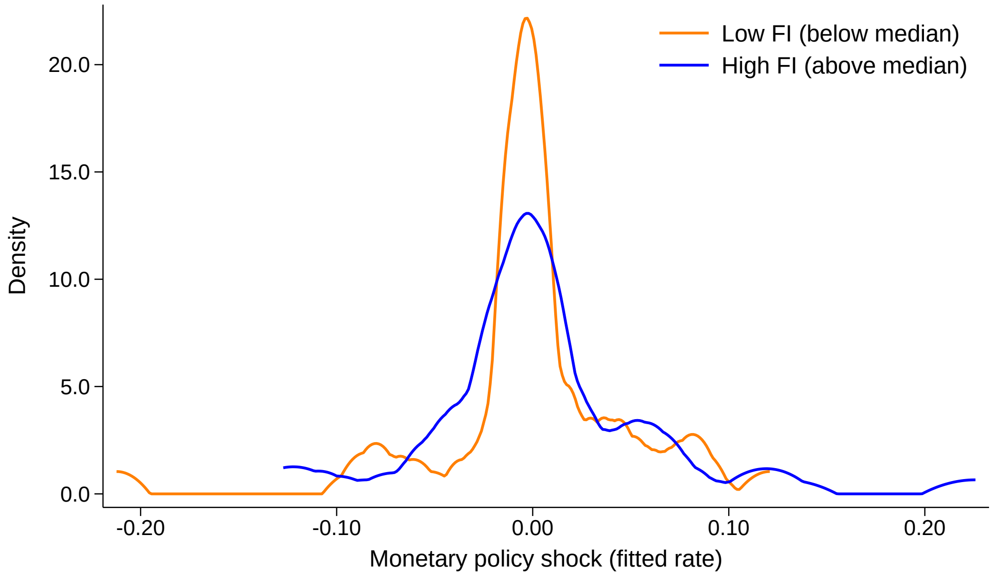
**Table B.6:** Tests of shock equality across FI regimes

| Test | Statistic | $p$-value | Conclusion |
|---|---|---|---|
| Two-sample $t$-test (equal means) | $t = -0.465$ | 0.643 | Cannot reject |
| Variance ratio $F$-test (equal variances) | $F = 0.771$ | 0.408 | Cannot reject |
| Kolmogorov–Smirnov (equal distributions) | $D = 0.095$ | 0.991 | Cannot reject |
| Wilcoxon rank-sum (equal distributions) | $z = 0.174$ | 0.862 | Cannot reject |

We further verify that the absolute magnitude of shocks does not differ across regimes: the mean of $|\widehat{\Delta i_t}|$ is 0.031 in low-FI quarters and 0.038 in high-FI quarters ($t = -0.708$, $p = 0.481$). The Pearson correlation between the shock and FI is $\rho = 0.008$ ($p = 0.940$); the correlation between the absolute shock and FI is $\rho = 0.017$ ($p = 0.881$). The monetary policy shock is therefore distributionally indistinguishable across FI regimes, ruling out the possibility that the interaction effect $\delta_h$ reflects the ECB delivering systematically different shocks during periods of high or low financial integration.

Figure B.4 provides a visual summary. The kernel density estimates of the shock distribution under low- and high-FI quarters are nearly indistinguishable: both are centered near zero with similar spread and tail behavior.

**Figure B.4:** Kernel density of monetary policy shocks by financial integration regime




*Notes:* Kernel density estimates of the quarterly monetary policy shock $\widehat{\Delta i_t}$ for low-FI (below median, orange) and high-FI (above median, blue) quarters. The sample is collapsed to 84 unique quarters (Q1 1999–Q4 2019), with 42 quarters per regime. The Kolmogorov–Smirnov test fails to reject distributional equality ($D = 0.095$, $p = 0.991$).

### B.1.5 Nonlinearity test: quadratic financial integration interaction

The baseline specification assumes that the interaction between FI and the monetary shock is linear: $\delta_h$ is constant across FI levels, so the marginal amplification effect of a unit increase in FI is the same regardless of the starting level. If the true relationship is nonlinear, the linear interaction may be misspecified. Following Tenreyro and Thwaites (2016), we test for nonlinearity by augmenting the specification with a quadratic interaction term $\widehat{\Delta i_t} \times FI_{t-1}^2$:

$$y_{i,t+h} - y_{i,t-1} = \alpha_{i,h} + \beta_h \widehat{\Delta i_t} + \delta_h \left(\widehat{\Delta i_t} \times FI_{t-1}\right) + \delta_h^2 \left(\widehat{\Delta i_t} \times FI_{t-1}^2\right) + \zeta_h FI_{t-1} + \phi_h FI_{t-1}^2 + \Gamma_h' X_{i,t} + \varepsilon_{i,t+h} \tag{B.2}$$

Both $FI_{t-1}$ and $FI_{t-1}^2$ are included as level controls. If $\delta_h^2$ is statistically insignificant, the linear specification is not rejected. A caveat is that the correlation between FI and FI$^2$ is 0.979, so the decomposition into linear and quadratic components is imprecise even when the joint effect is well identified.

**HICP: linear specification not rejected.** Table B.7 reports $\delta_h^2$ for HICP at all 17 horizons. The quadratic term is insignificant throughout—the largest $|t|$-statistic is 1.16 (at $h = 10$)—and 0 out of 17 horizons are significant at the 90% confidence level. The linear interaction adequately captures the relationship between FI and price transmission.

**Table B.7:** Quadratic interaction coefficient $\delta_h^2$ — HICP

| $h$ | $\delta_h^2$ (SE) | $t$ | $h$ | $\delta_h^2$ (SE) | $t$ |
|---|---|---|---|---|---|
| 0 | −29 (140) | −0.20 | 9 | −537 (486) | −1.11 |
| 1 | −134 (180) | −0.74 | 10 | −586 (506) | −1.16 |
| 2 | −126 (261) | −0.48 | 11 | −497 (514) | −0.97 |
| 3 | −258 (287) | −0.90 | 12 | −328 (566) | −0.58 |
| 4 | −108 (361) | −0.30 | 13 | −257 (562) | −0.46 |
| 5 | −256 (429) | −0.60 | 14 | −298 (598) | −0.50 |
| 6 | −440 (478) | −0.92 | 15 | −468 (544) | −0.86 |
| 7 | −482 (482) | −1.00 | 16 | −272 (532) | −0.51 |
| 8 | −442 (517) | −0.85 | | | |
| Significant at 90% confidence: 0/17. | | | | | |

*Notes:* Driscoll–Kraay standard errors with 4 lags. The quadratic interaction is $\widehat{\Delta i_t} \times FI_{t-1}^2$.

**GDP: nonlinearity that strengthens the finding.** Table B.8 shows that the quadratic term is negative and significant at 13 out of 17 horizons for GDP, concentrated at $h = 0$, $h = 2$–10, and $h = 14$–16. The negative $\delta_h^2$ means the coefficient polynomial $(\beta_h + \delta_h FI + \delta_h^2 FI^2)$ is concave, which—after the sign flip to expansionary IRFs—implies that the marginal amplification effect of FI on the reported IRF increases at higher levels of integration.

**Table B.8:** Quadratic interaction coefficient $\delta_h^2$ — real GDP

| $h$ | $\delta_h^2$ (SE) | $t$ | $h$ | $\delta_h^2$ (SE) | $t$ |
|---|---|---|---|---|---|
| 0 | −300 (174) | −1.73 | 9 | −1206** (454) | −2.66 |
| 1 | −638 (407) | −1.57 | 10 | −1013** (482) | −2.10 |
| 2 | −823* (453) | −1.82 | 11 | −632 (519) | −1.22 |
| 3 | −918** (434) | −2.12 | 12 | −712 (514) | −1.39 |
| 4 | −1082** (466) | −2.32 | 13 | −776* (476) | −1.63 |
| 5 | −1084** (453) | −2.39 | 14 | −981** (484) | −2.02 |
| 6 | −1142** (473) | −2.41 | 15 | −1258*** (491) | −2.56 |
| 7 | −1290*** (464) | −2.78 | 16 | −1594*** (604) | −2.64 |
| 8 | −1174*** (444) | −2.65 | | | |
| Significant at 90% confidence: 13/17. | | | | | |

*Notes:* See notes to Table B.7. Dependent variable is log real GDP.

The nonlinearity leaves the economic conclusion intact. Table B.9 compares the implied IRF differences between the 75th and 25th percentiles of FI under the linear and quadratic specifications. For HICP, the quadratic model produces regime differences that are 10–20% smaller at peak horizons, but the qualitative pattern is identical, indicating that the linear model is not a material misspecification for price transmission. For GDP, the quadratic model produces *larger* regime differences at the short-to-medium horizons ($h = 0$–10) where the amplification effect is economically meaningful—the nonlinearity implies that the marginal amplification effect of FI strengthens at higher levels of integration. If anything, the linear specification understates the amplification of GDP responses.

**Table B.9:** Implied IRF differences (p75 − p25) under linear vs. quadratic specification

| | HICP | | Real GDP | |
|---|---|---|---|---|
| $h$ | Linear | Quadratic | Linear | Quadratic |
| 0 | 3.3 | 0.8 | 20.7 | 24.8 |
| 1 | 14.1 | 11.2 | 47.7 | 60.6 |
| 2 | 27.9 | 24.0 | 48.6 | 65.4 |
| 4 | 43.0 | 31.0 | 81.3 | 104.9 |
| 5 | 57.2 | 44.0 | 77.6 | 100.8 |
| 8 | 77.9 | 60.2 | 52.2 | 100.3 |
| 10 | 82.4 | 70.0 | 29.1 | 101.2 |
| 12 | 76.5 | 64.3 | −7.8 | 80.1 |

*Notes:* Units: $\times 10^{-2}$ percentage points. Values computed as $\text{IRF}(h, FI_{p75}) - \text{IRF}(h, FI_{p25})$, where under the linear model $\text{IRF}(h, FI) = -(\beta_h + \delta_h FI) \times \sigma_{\widehat{\Delta i}} \times 100$ and under the quadratic model $\text{IRF}(h, FI) = -(\beta_h + \delta_h FI + \delta_h^2 FI^2) \times \sigma_{\widehat{\Delta i}} \times 100$.

### B.1.6 Multiple threshold robustness

The paper reports impulse responses evaluated at the 25th and 75th percentiles of FI. Because FI enters Equation 2 as a continuous variable, the IRF at any FI level is an analytical function of the estimated coefficients:

$$\text{IRF}(h, FI) = -(\beta_h + \delta_h \times FI) \times \sigma_{\widehat{\Delta i}} \times 100 \tag{B.3}$$

No re-estimation is needed to evaluate the IRF at different FI values. The regime difference for a threshold pair $(FI_{\text{low}}, FI_{\text{high}})$ is $-\delta_h \times (FI_{\text{high}} - FI_{\text{low}}) \times \sigma_{\widehat{\Delta i}} \times 100$, which scales linearly with the FI gap.

We evaluate the IRF at four threshold pairs spanning a range of widths: p20/p80 (*FI* gap = 0.150), p25/p75 (*FI* gap = 0.135, baseline), p30/p70 (*FI* gap = 0.098), and p33/p67 (*FI* gap = 0.078). Table B.10 reports the peak regime differences.

**Table B.10:** Peak regime differences across threshold pairs

| Threshold pair | FI gap | HICP peak | at $h$ | GDP peak | at $h$ | Ratio to baseline |
|---|---|---|---|---|---|---|
| p20/p80 | 0.1501 | 95.6 | 9 | 90.4 | 4 | 1.111 |
| p25/p75 (baseline) | 0.1351 | **86.0** | **9** | **81.3** | **4** | 1.000 |
| p30/p70 | 0.0980 | 62.4 | 9 | 59.0 | 4 | 0.726 |
| p33/p67 | 0.0779 | 49.6 | 9 | 46.9 | 4 | 0.577 |

*Notes:* Units: $\times 10^{-2}$ percentage points (raw, unsmoothed). The ratio column reports the ratio of each pair's peak HICP difference to the baseline p25/p75 difference; the ratios are identical for GDP and match the ratio of FI gaps exactly.

Three features confirm the robustness of the threshold choice. First, the peak horizon is identical across all four pairs: $h = 9$ for HICP and $h = 4$ for GDP ($h = 10$ and $h = 4$, respectively, after smoothing). The timing of maximum amplification does not depend on the threshold. Second, the regime differences are strictly monotonically ordered at every horizon for HICP (all 17 horizons satisfy $\text{Diff}_{p20/p80} > \text{Diff}_{p25/p75} > \text{Diff}_{p30/p70} > \text{Diff}_{p33/p67}$). For GDP, 12 out of 17 horizons satisfy this ordering, with the 5 exceptions occurring at $h$ = 12–16 where $\delta_h$ changes sign. Third, the ratios of peak differences across threshold pairs match the ratios of FI gaps to four decimal places—a mechanical consequence of the linear interaction that confirms the internal consistency of the continuous specification. A reader who prefers terciles (p33/p67) would see the same amplification pattern at 58% of the quartile-based magnitude; a reader who prefers more extreme thresholds (p20/p80) would see 111%.

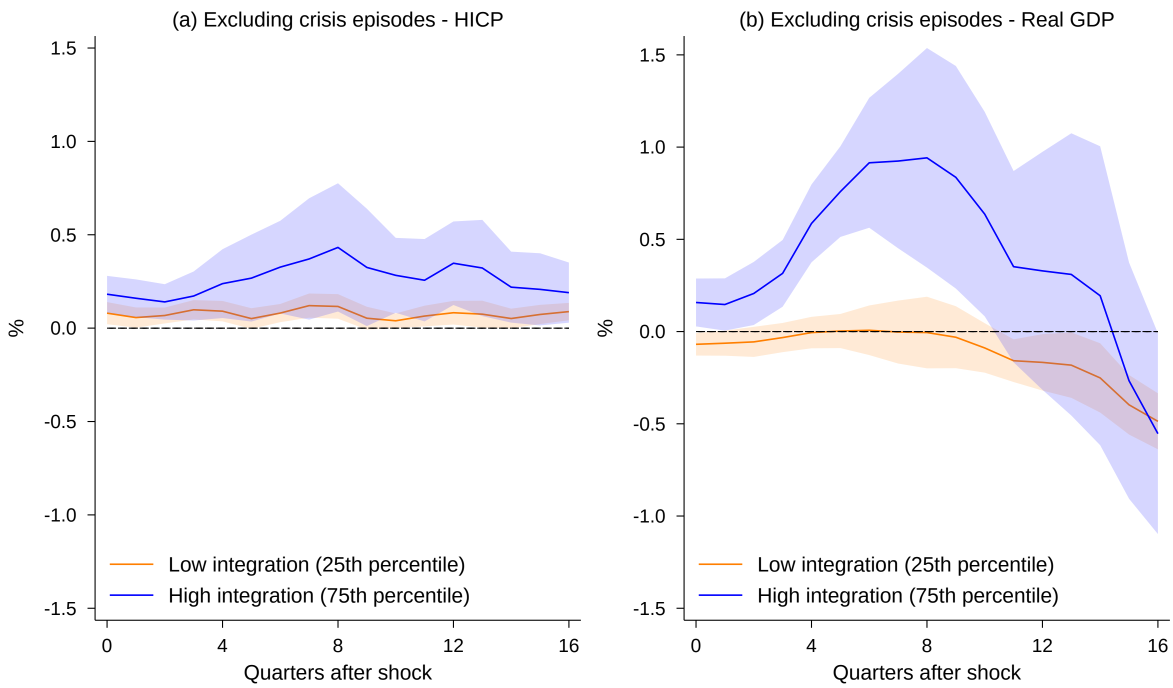
## B.2 Sensitivity to Data and Sample Choices

### B.2.1 Excluding crisis episodes

As a complement to the recession-indicator approach in Appendix B.1.2, we address cyclical confounding more aggressively by restricting the sample to end in Q2 2007, excluding the global financial crisis and the EA sovereign debt crisis entirely. This removes periods of severe macro-financial stress where financial integration and policy transmission are most likely to co-move with the business cycle. The shorter sample necessarily involves a substantial reduction in observations and statistical power.

Figure B.5 shows that the amplification pattern is preserved: higher financial integration is associated with stronger monetary policy transmission to both prices and output, even outside crisis periods.

**Figure B.5:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Sample ends in Q2 2007 instead of Q4 2019. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

### B.2.2 Crisis down-weighting

The pre-2007 sample robustness test (Appendix B.2.1) removes all data after mid-2007, discarding both the crisis and the post-crisis recovery. A complementary approach retains the full sample but down-weights crisis-contaminated observations, applying the variance down-weighting methodology of Lenza and Primiceri (2022) to the crisis period (Q3 2007–Q2 2013), analogous to the COVID-period treatment in Appendix B.2.4.

**Crisis contamination in local projections.** An observation at quarter $t$ and horizon $h$ is crisis-contaminated if any quarter in the dependent-variable window $[t-1, t+h]$ overlaps with the crisis period (Q3 2007–Q2 2013). At $h = 0$, approximately 29% of observations are contaminated (those with $t \in [\text{Q2 2007}, \text{Q2 2013}]$); at $h = 16$, this rises because observations from as early as Q3 2003 have their $h = 16$ dependent variable reaching into the crisis window. Unlike a donut specification—which would excise crisis observations and break the autocorrelation structure of the panel—the down-weighting approach preserves all observations and lets the data determine how much to discount crisis-era variation.

**Two-pass estimation.** For each outcome and horizon $h$, we estimate a two-pass procedure identical to the COVID treatment: (i) run the baseline specification with a crisis-contamination dummy to obtain residuals and compute the variance ratio $\lambda_h = \text{Var}(\hat{e} \mid D_h^{crisis} = 1)/\text{Var}(\hat{e} \mid D_h^{crisis} = 0)$; (ii) re-estimate via WLS, assigning weight $\omega = 1/\lambda_h$ to contaminated observations and $\omega = 1$ to normal observations. Because `xtscc` does not support analytical weights, WLS is implemented via a manual within-transformation with Driscoll–Kraay standard errors, as in Appendix B.2.4.

**Variance ratios.** Table B.11 reports $\lambda_h$ for both outcome variables. For HICP, $\lambda_h$ peaks at 2.31 ($h = 3$) and decays to 1.15 by $h = 16$, reflecting the gradual nature of price adjustment even during crises. For GDP, $\lambda_h$ peaks at 2.14 ($h = 2$) and decays toward 1 at longer horizons ($\lambda_{15} = 1.01$), falling below 1 at $h = 16$ ($\lambda_{16} = 0.95$), where crisis-contaminated residuals are slightly *smaller* than normal. The WLS procedure therefore down-weights crisis observations most heavily at short-to-medium horizons where crisis distortion is largest.

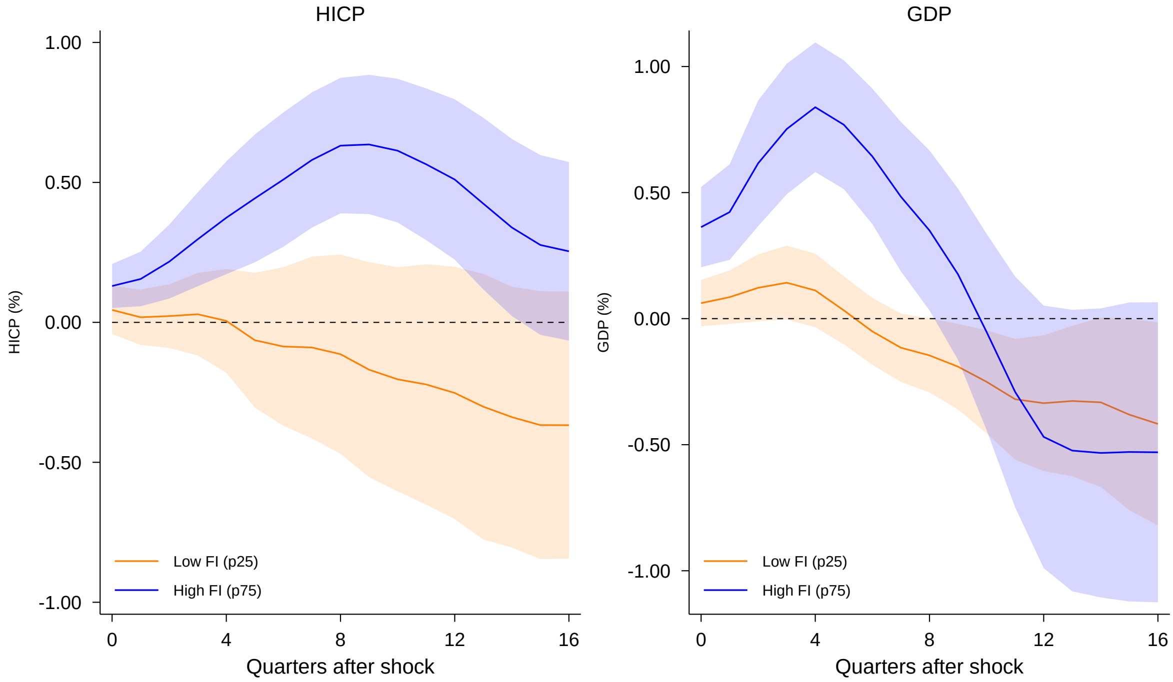
**Table B.11:** Variance ratio $\lambda_h$ (crisis / normal residual variance) by horizon

| $h$ | HICP | GDP | $h$ | HICP | GDP |
|---|---|---|---|---|---|
| 0 | 1.69 | 1.75 | 9 | 1.27 | 1.73 |
| 1 | 2.13 | 2.10 | 10 | 1.27 | 1.56 |
| 2 | 2.29 | 2.14 | 11 | 1.23 | 1.40 |
| 3 | 2.31 | 2.09 | 12 | 1.19 | 1.25 |
| 4 | 2.13 | 2.01 | 13 | 1.23 | 1.15 |
| 5 | 2.02 | 1.87 | 14 | 1.23 | 1.07 |
| 6 | 1.81 | 1.85 | 15 | 1.18 | 1.01 |
| 7 | 1.57 | 1.84 | 16 | 1.15 | 0.95 |
| 8 | 1.36 | 1.81 | | | |

*Notes:* $\lambda_h > 1$ indicates that crisis-contaminated observations have larger residual variance than normal observations at horizon $h$. Weights are $\omega = 1/\lambda_h$ for contaminated observations and $\omega = 1$ for normal observations.

**Results.** Figure B.6 reports the impulse responses under crisis down-weighting. The amplification pattern is preserved: higher financial integration leads to substantially larger macroeconomic responses. For HICP, the FI interaction coefficient $\delta_h$ is negative and statistically significant ($|t| > 2$) at 11 of 17 horizons ($h = 2$, $h = 4$–13), with $t$-statistics reaching $-2.97$ at the peak ($h = 9$). For GDP, $\delta_h$ is significant at $h = 0$–7 (8 of 17 horizons), peaking at $h = 4$ ($t = -3.70$), and fading at longer horizons as expected from the baseline. The amplification of monetary transmission by financial integration persists after crisis down-weighting, indicating that the baseline finding is not driven primarily by crisis-period volatility.

**Figure B.6:** Crisis down-weighting: responses of HICP and real GDP under low and high financial integration




*Notes:* Crisis-period observations (Q3 2007–Q2 2013) are down-weighted using the Lenza and Primiceri (2022) variance-ratio approach. Impulse responses (in %) to a one-standard-deviation expansionary monetary policy shock, evaluated at the 25th percentile (orange) and 75th percentile (blue) of the FI distribution. Shaded areas denote 68% confidence intervals based on Driscoll-Kraay standard errors.

### B.2.3 Higher frequency data

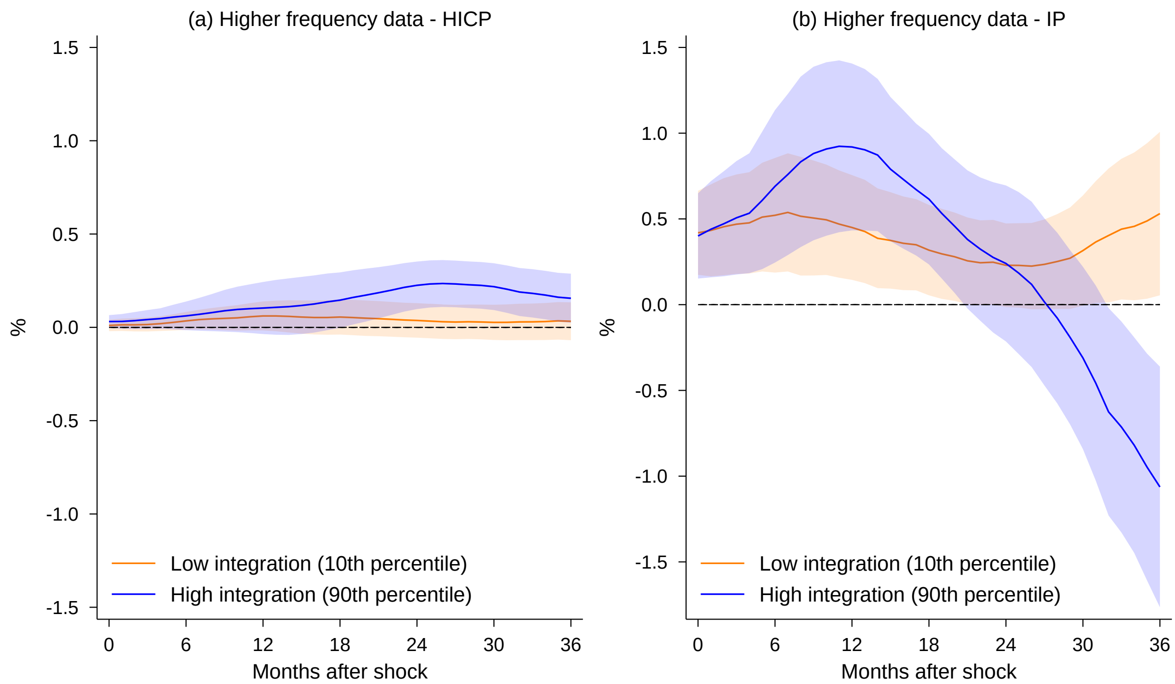
Recent literature shows that the time aggregation of economic activity and monetary policy shocks can influence the identification of monetary policy transmission (Buda et al., 2025). In particular, when high-frequency shocks are aggregated to match the lower frequency of standard macroeconomic variables, such as quarterly real GDP, the empirical response can be shifted to longer lags.

To address this concern, we estimate the impulse responses using higher-frequency data. Specifically, we use monthly observations for HICP and Industrial Production (IP) from Eurostat. IP, similarly to HICP, is set to 2015=100. Also, IP is seasonally and calendar adjusted, while for HICP we seasonally adjust the data using JDemetra+. The monetary policy shocks are still from Jarociński and Karadi (2020) but in monthly frequency to match these variables. For financial integration, we rely on the price-based indicator from Hoffmann, Kremer and Zaharia (2020), which is available at monthly frequency. In terms of the specification used to estimate the impulse responses, we adjust to increase the number of lags of the dependent variable from 4 to 12.

Note that this exercise simultaneously changes three features of the estimation relative to the baseline: the data frequency (monthly instead of quarterly), the output measure (IP instead of GDP), and the financial integration indicator (price-based instead of quantity-based). Appendix B.3.4 varies the FI indicator at quarterly frequency while holding the other features fixed, partially disentangling these channels.

Figure B.7 reports the impulse responses under low (10th percentile) and high (90th percentile) levels of financial integration. For HICP, the response under high integration exceeds the low-integration response from month 6 onwards. For IP, the high-integration response peaks at approximately 0.9% around month 12, compared to roughly 0.5% under low integration. The confidence bands overlap at several horizons, so this exercise provides suggestive rather than definitive evidence that the amplification extends to a monthly specification with a price-based integration measure. Appendix B.3.4 shows that using the price-based indicator at quarterly frequency preserves the amplification more clearly, suggesting that neither time aggregation nor the specific integration indicator alone drives the baseline conclusions.

**Figure B.7:** Responses of HICP and IP to a monetary policy shock under low and high levels of financial integration (monthly data)




*Notes:* Monthly data is employed for all variables in Equation 2. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (10th percentile, orange line) and high (90th percentile, blue line) levels of financial integration. The financial integration indicator is the price-based measure from Hoffmann, Kremer and Zaharia (2020). Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors. Coefficients are smoothed using a centered moving average with a 9-month window to match the time coverage of the quarterly baseline's 3-quarter smoothing.

#### B.2.4 Extended time frame with Lenza–Primiceri COVID handling

The baseline sample ends in Q4 2019, excluding the COVID-19 pandemic entirely. We extend the sample through Q1 2023—the last quarter for which the financial integration indicator is available—and address the pandemic period using the variance down-weighting approach of Lenza and Primiceri (2022), adapted to the local projection setting.

**COVID contamination in local projections.** In the LP framework, the dependent variable $y_{i,t+h} - y_{i,t-1}$ spans the window $[t-1, t+h]$. An observation is COVID-contaminated at horizon $h$ if *any* quarter in this window falls within the acute pandemic period (Q1 2020–Q2 2020). This creates horizon-specific contamination: at $h = 0$, approximately 3% of observations are contaminated (57 of 1,843 country-quarters); at $h = 16$, this rises to 20% (361 of 1,843), because the forward-looking structure of LPs means that observations from as early as Q1 2016 have their $h = 16$ dependent variable touching the COVID window.

**Two-pass estimation.** For each outcome variable and each horizon $h$, we implement a two-pass procedure:

1. **Pass 1 — Variance ratio estimation.** We run the baseline specification (Equation 2) augmented with a contamination dummy $D_h^{covid}$ and extract the residuals $\hat{e}_{i,t,h}$. The variance ratio is $\lambda_h = \text{Var}(\hat{e}|D_h^{covid} = 1)/\text{Var}(\hat{e}|D_h^{covid} = 0)$.
2. **Pass 2 — Weighted least squares.** We assign weight $\omega = 1$ to normal observations and $\omega = 1/\lambda_h$ to contaminated observations, then estimate via WLS. Because the contamination status depends only on the time dimension—not on the country—the weights $\omega$ are identical across countries within each quarter. Because `xtscc` does not support analytical weights, we implement WLS via a manual within-transformation: we demean all variables within country (replicating the fixed-effects transformation), multiply all demeaned variables by $\sqrt{\omega}$, and estimate the transformed regression with Driscoll–Kraay standard errors. The time-only variation in weights ensures that this unweighted-then-weighted procedure closely approximates the weighted fixed-effects estimator.

**Variance ratios.** Table B.12 reports $\lambda_h$ for both outcome variables. For HICP, the variance ratio is below 1 at short horizons (COVID-contaminated residuals are *smaller* than normal, because the price level was relatively stable in Q1–Q2 2020), which under the $\omega = 1/\lambda_h$ rule implies that these observations are *up-weighted* rather than down-weighted at short horizons—consistent with their lower residual variance, as pure inverse-variance weighting dictates. The ratio rises above 1 at medium horizons ($\lambda \approx 2.4$ at $h = 10$), where the procedure down-weights as expected. For GDP, $\lambda_h$ is very large at impact ($\lambda = 23.07$ at $h = 0$), reflecting the extreme GDP collapse in Q1–Q2 2020, and decays to approximately 1 by $h = 10$ ($\lambda_{10} = 1.03$). The WLS procedure therefore heavily down-weights the COVID observations for GDP at short horizons—exactly where the pandemic distortion is most severe—while leaving them nearly unweighted at medium horizons. At long horizons ($h \geqslant 11$), $\lambda_h$ falls slightly below 1 for GDP, implying mild up-weighting analogous to the HICP pattern at short horizons.

**Table B.12:** Variance ratio $\lambda_h$ (COVID / normal residual variance) by horizon

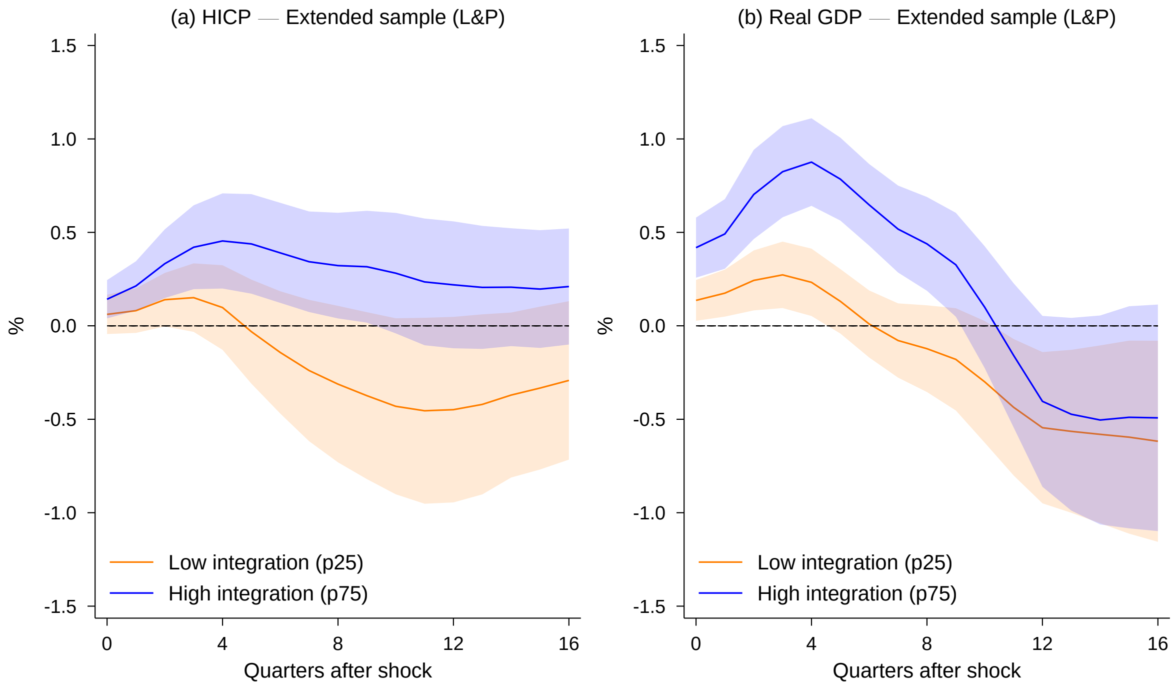
| $h$ | HICP | GDP | $h$ | HICP | GDP |
|---|---|---|---|---|---|
| 0 | 0.80 | 23.07 | 9 | 2.00 | 1.20 |
| 1 | 0.32 | 10.46 | 10 | 2.40 | 1.03 |
| 2 | 0.28 | 7.47 | 11 | 1.97 | 0.88 |
| 3 | 0.41 | 5.13 | 12 | 1.82 | 0.86 |
| 4 | 0.35 | 3.94 | 13 | 1.70 | 0.89 |
| 5 | 0.40 | 2.98 | 14 | 1.61 | 0.88 |
| 6 | 0.55 | 2.30 | 15 | 1.58 | 0.90 |
| 7 | 1.03 | 1.70 | 16 | 1.41 | 0.84 |
| 8 | 1.51 | 1.42 | | | |

*Notes:* $\lambda_h > 1$ indicates that COVID-contaminated observations have larger residual variance than normal observations at horizon $h$. Weights are $\omega = 1/\lambda_h$ for contaminated observations and $\omega = 1$ for normal observations.

**Extended sample parameters.** The first stage uses all available ECB meetings through Q2 2023 ($N = 292$, $F = 86.49$; fitted shocks through Q1 2023—the last quarter for which FI data are available—enter the second-stage LP, compared to 265 meetings and $F = 79.95$ in the baseline). The extended-sample FI percentiles are: p25 = 0.253, p75 = 0.369, SD = 0.082 (baseline: p25 = 0.242, p75 = 0.377, SD = 0.087). The SD of the fitted shock is 0.057 (baseline: 0.056).

**Results.** Figure B.8 shows the impulse responses under the extended sample with L&P weighting. The amplification pattern is preserved: higher financial integration leads to substantially larger macroeconomic responses at all horizons through $h = 10$. The regime difference (p75 minus p25) peaks at 0.71 p.p. for HICP (baseline: 0.83 p.p., ratio 0.85) and 0.66 p.p. for GDP (baseline: 0.78 p.p., ratio 0.85).

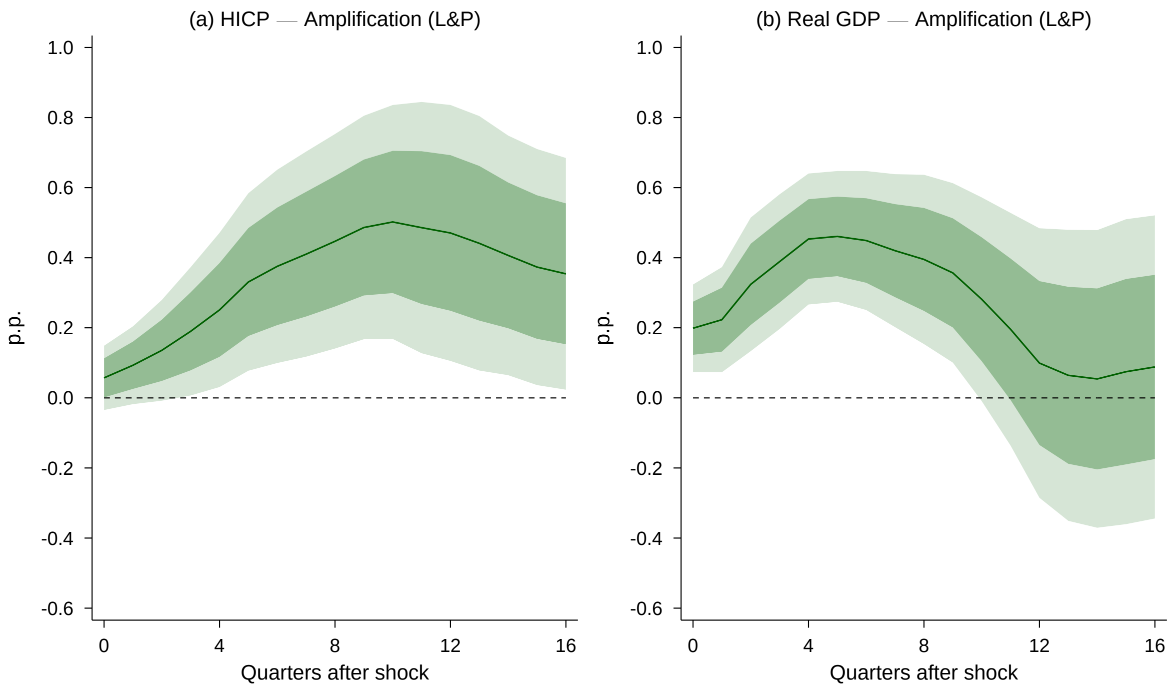
**Figure B.8:** Extended sample (Q1 1999–Q1 2023) with Lenza–Primiceri COVID handling: responses of HICP and real GDP under low and high financial integration




*Notes:* Sample extended to Q1 2023. COVID observations (Q1–Q2 2020) are down-weighted using the Lenza and Primiceri (2022) variance-ratio approach. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

Figure B.9 reports the amplification (marginal effect of a 1-SD increase in FI on the IRF). The peak HICP amplification is 0.50 p.p. at $h = 10$ (baseline: 0.53 p.p. at $h = 10$; ratio 0.94). The peak GDP amplification is 0.47 p.p. at $h = 5$ (baseline: 0.50 p.p. at $h = 4$; ratio 0.93). Both peaks are statistically significant at the 90% confidence level.

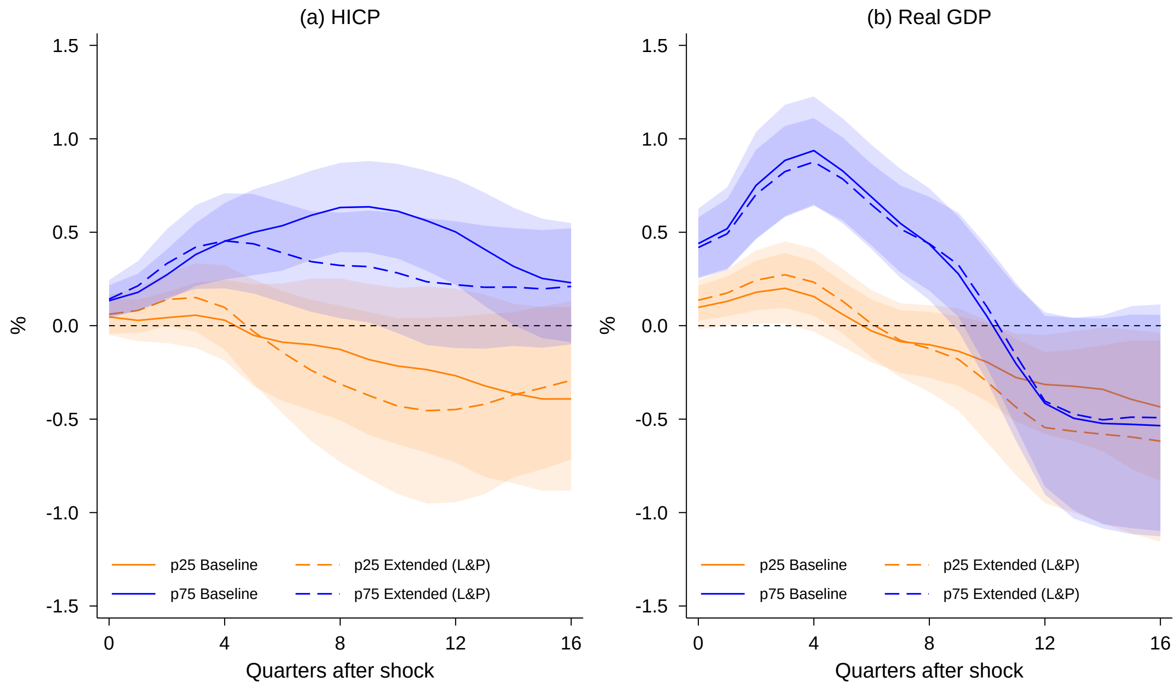
**Figure B.9:** Extended sample with L&P handling: amplification of monetary policy by financial integration




*Notes:* Amplification effect: change in the impulse response for a one-standard-deviation increase in financial integration. Sample extended to Q1 2023 with Lenza and Primiceri (2022) COVID handling. Dark shaded areas denote 68% confidence intervals; light shaded areas denote 90% confidence intervals.

Figure B.10 overlays the baseline (Q1 1999–Q4 2019) and extended-sample IRFs for direct comparison. The solid lines (baseline) and dashed lines (extended) track each other closely, indicating that the amplification result is not sensitive to the sample endpoint or the treatment of the pandemic period.

**Figure B.10:** Comparison of baseline and extended-sample impulse responses




*Notes:* Solid lines: baseline (Q1 1999–Q4 2019). Dashed lines: extended sample (Q1 1999–Q1 2023) with Lenza and Primiceri (2022) COVID handling. Orange: 25th percentile of FI. Blue: 75th percentile of FI. Shaded areas denote 68% confidence intervals.

Table B.13 summarizes the comparison. The extended-sample estimates retain 93–94% of the baseline amplification magnitudes, with identical peak horizons for HICP and a one-quarter shift for GDP. The regime differences retain 85% of the baseline. The amplification of monetary transmission by financial integration therefore extends to a sample that includes the pandemic, handled via a principled variance down-weighting approach.

**Table B.13:** Comparison of baseline and extended-sample results

| | Peak amplification (p.p.) | | Peak regime difference (p.p.) | |
|---|---|---|---|---|
| | Baseline | Extended (L&P) | Baseline | Extended (L&P) |
| HICP | 0.53 ($h = 10$) | 0.50 ($h = 10$) | 0.83 ($h = 10$) | 0.71 ($h = 10$) |
| GDP | 0.50 ($h = 4$) | 0.47 ($h = 5$) | 0.78 ($h = 4$) | 0.66 ($h = 5$) |
| Ratio (Ext/Base) | 0.94 (HICP), 0.93 (GDP) | | 0.85 (HICP), 0.85 (GDP) | |

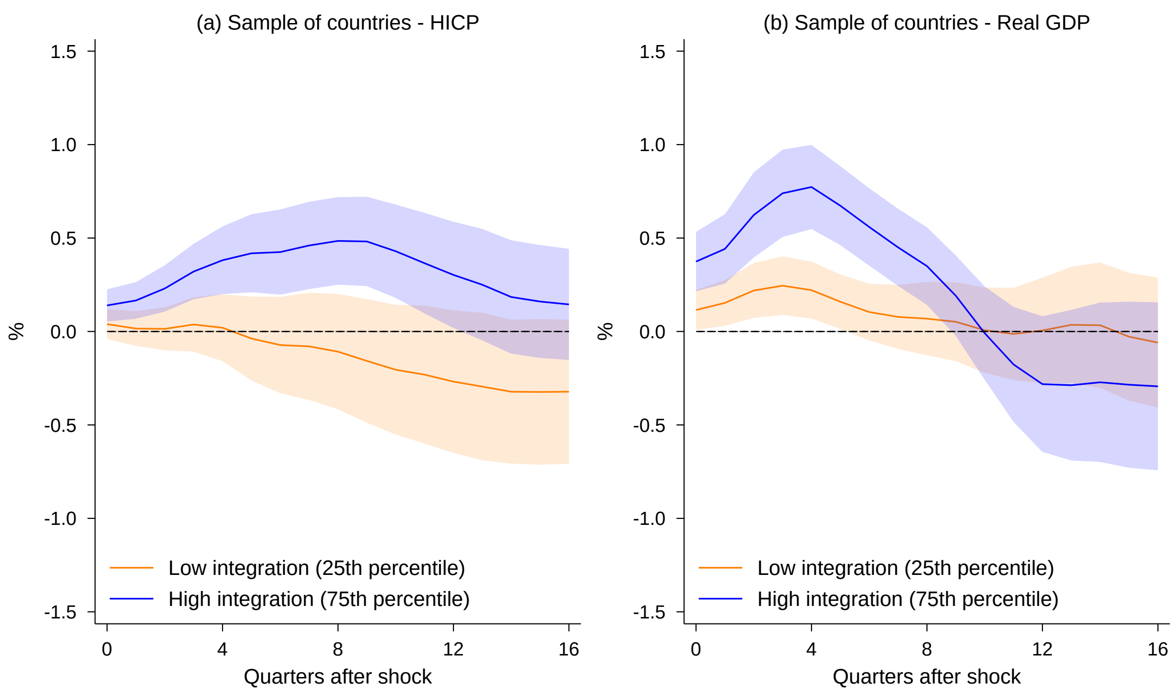
*Notes:* Amplification is the change in the smoothed impulse response for a 1-SD increase in FI. Regime difference is the smoothed IRF at the 75th percentile minus the 25th percentile of FI.

### B.2.5 Sample of countries

To check if our findings hold with a different sample, we analyze only the eleven original EA countries that adopted the euro in 1999: Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal, and Spain. Our baseline analysis includes all EA countries except Croatia, which joined in 2023. The financial integration indicator, described in Section 2.1, is based mainly on data from these eleven countries. By focusing on them, we test whether including later EA members, which may have different levels of financial integration or economic traits, affects our results. We use the same specification as in Equation 2.

Figure B.11 shows that the amplification pattern is preserved in this smaller sample. Higher financial integration is associated with stronger effects of monetary policy shocks on prices and production, suggesting that including later EA members does not materially alter the baseline findings.

**Figure B.11:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration



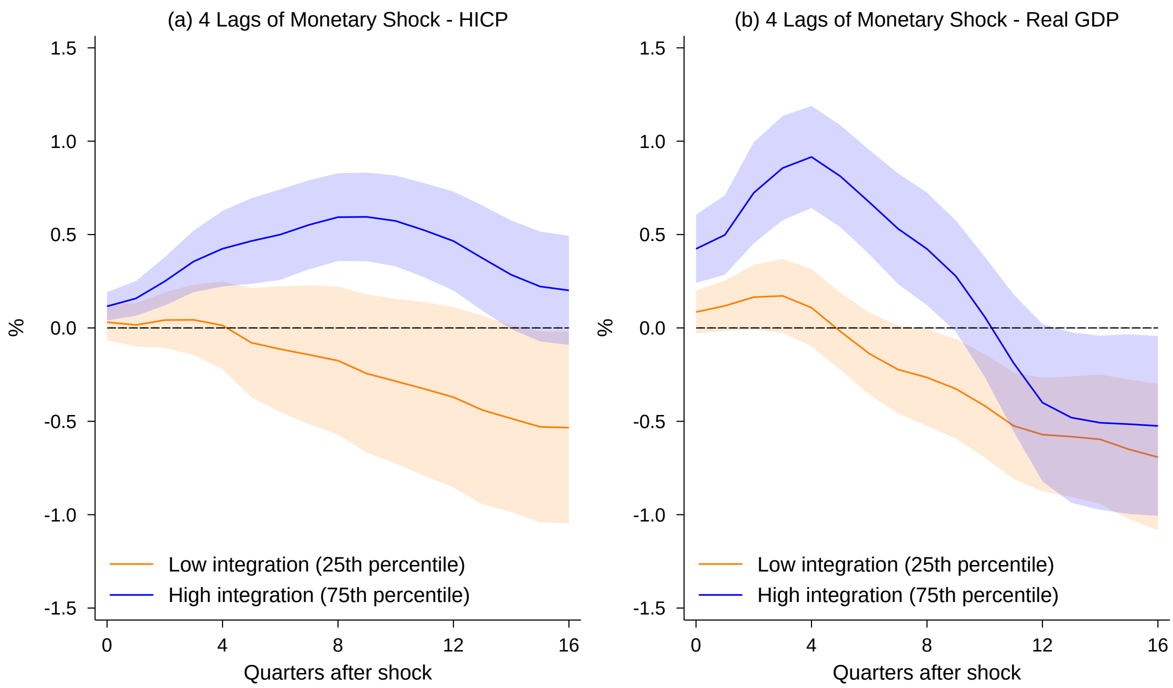

*Notes:* Sample of countries includes the original eleven member states of the EA instead of nineteen. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

## B.3 Alternative Specifications and Shock Measures

### B.3.1 Number of lags of monetary policy shocks and dependent variable

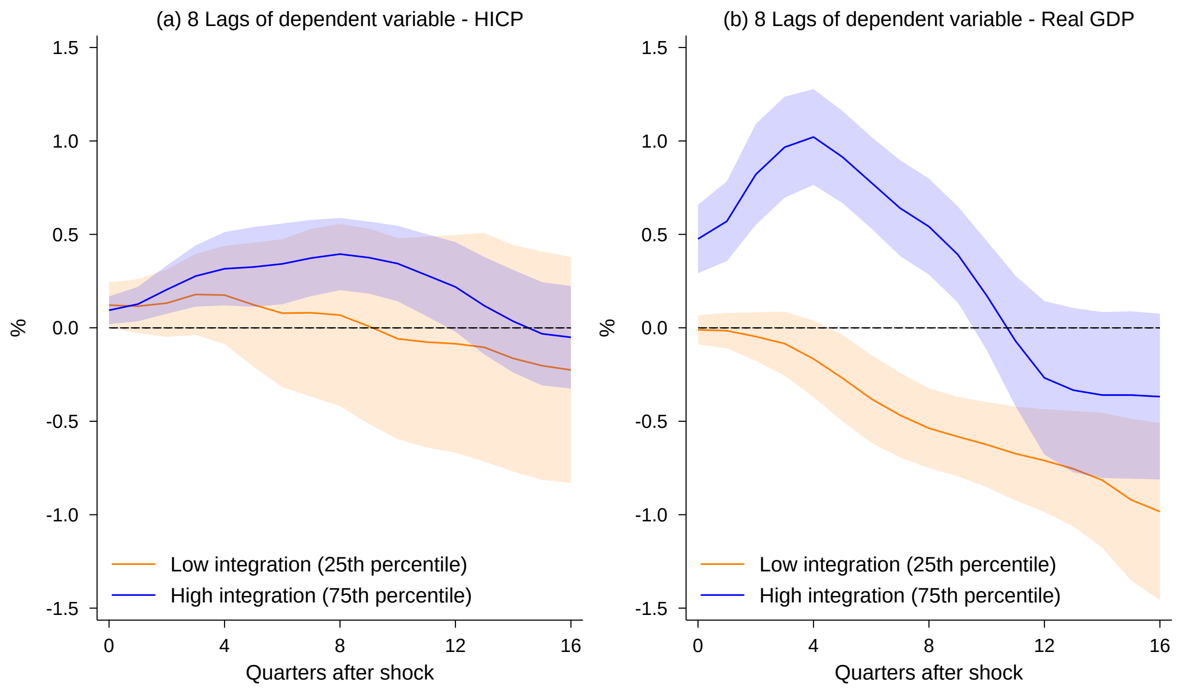
We test sensitivity to the lag structure by estimating two alternative specifications: increasing the monetary shock lags from one to four (Figure B.12), and extending the dependent-variable lags from four to eight (Figure B.13). Under both alternatives, the amplification pattern is unchanged: responses remain larger under high financial integration at the same horizons as in the baseline.

**Figure B.12:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Specification in Equation 2 has four instead of one lag of monetary shock. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

**Figure B.13:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration

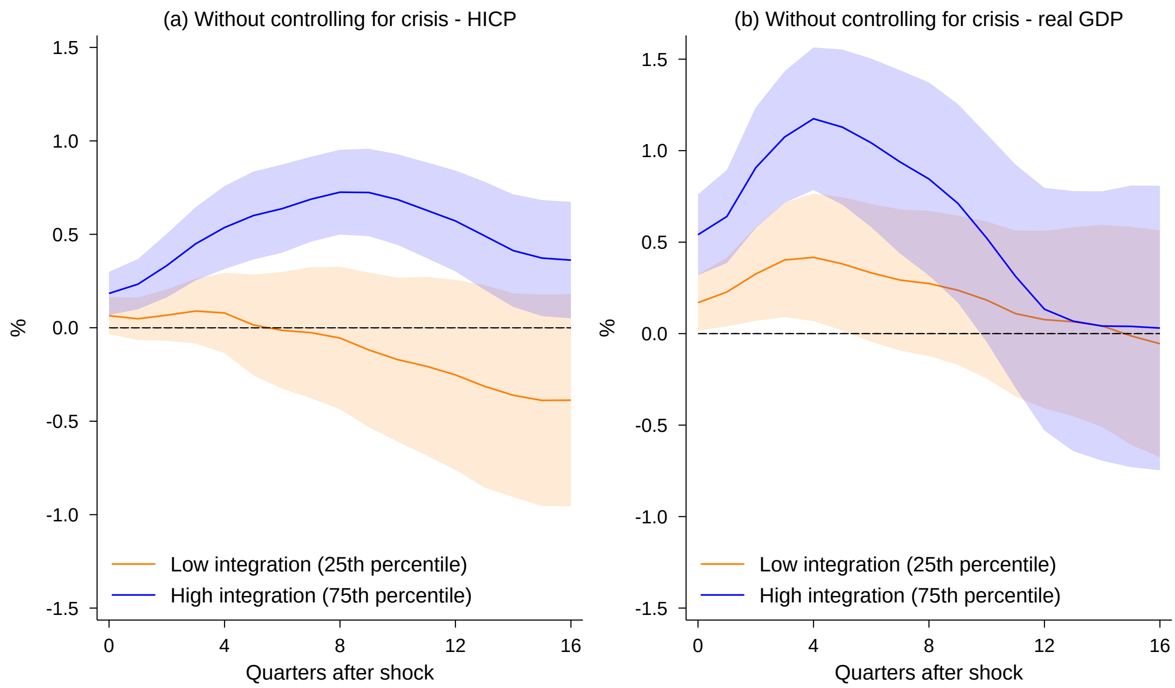



*Notes:* Specification in Equation 2 has eight instead of four lags of dependent variable. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

### B.3.2 Without controlling for crisis

We re-estimate Equation 2 after removing the three crisis-period dummies, which absorb common time-varying shocks during the GFC and sovereign debt crisis. Figure B.14 shows that the impulse responses are quantitatively close to the baseline, with the amplification pattern preserved at all horizons. This supports the interpretation that the baseline results are not driven by a few large common shocks captured by the time dummies.

**Figure B.14:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration

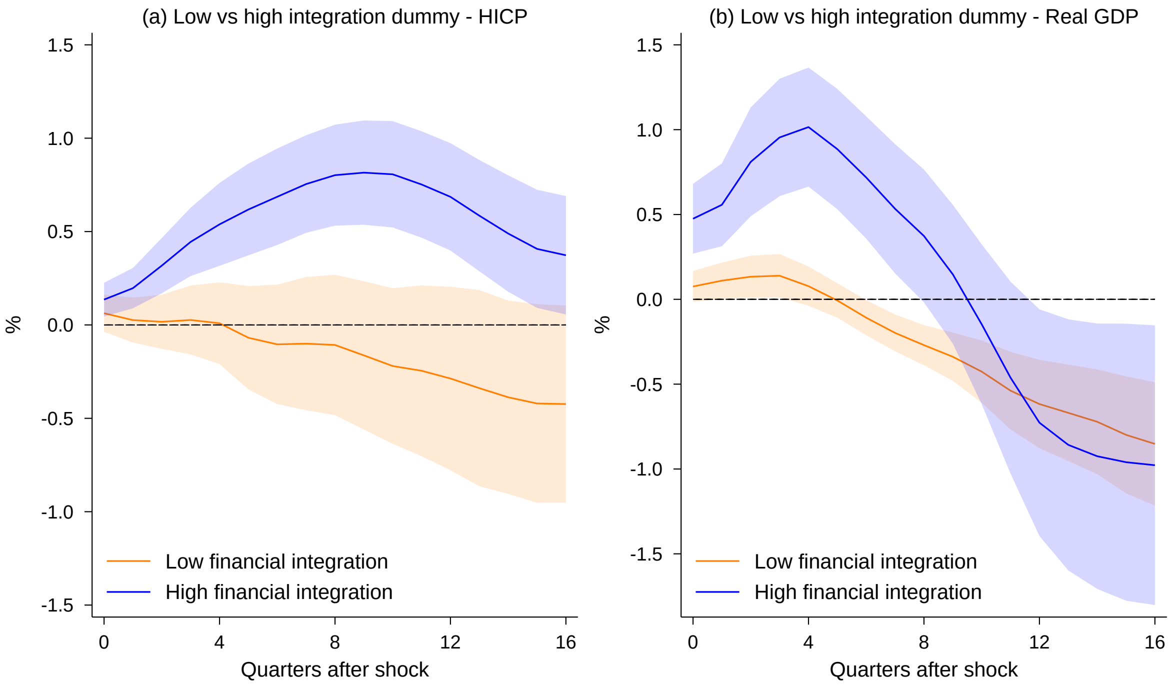



*Notes:* The dummies controlling for the crisis period are removed from the specification. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

### B.3.3 Financial integration regimes using a dummy

We replace the continuous FI measure in Equation 2 with a binary indicator equal to one when the composite indicator exceeds its median value (0.2988). This tests whether the amplification depends on the continuous functional form. Figure B.15 shows that the pattern holds: responses to a monetary shock are stronger during high-integration periods, consistent with the continuous-measure baseline.

**Figure B.15:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration



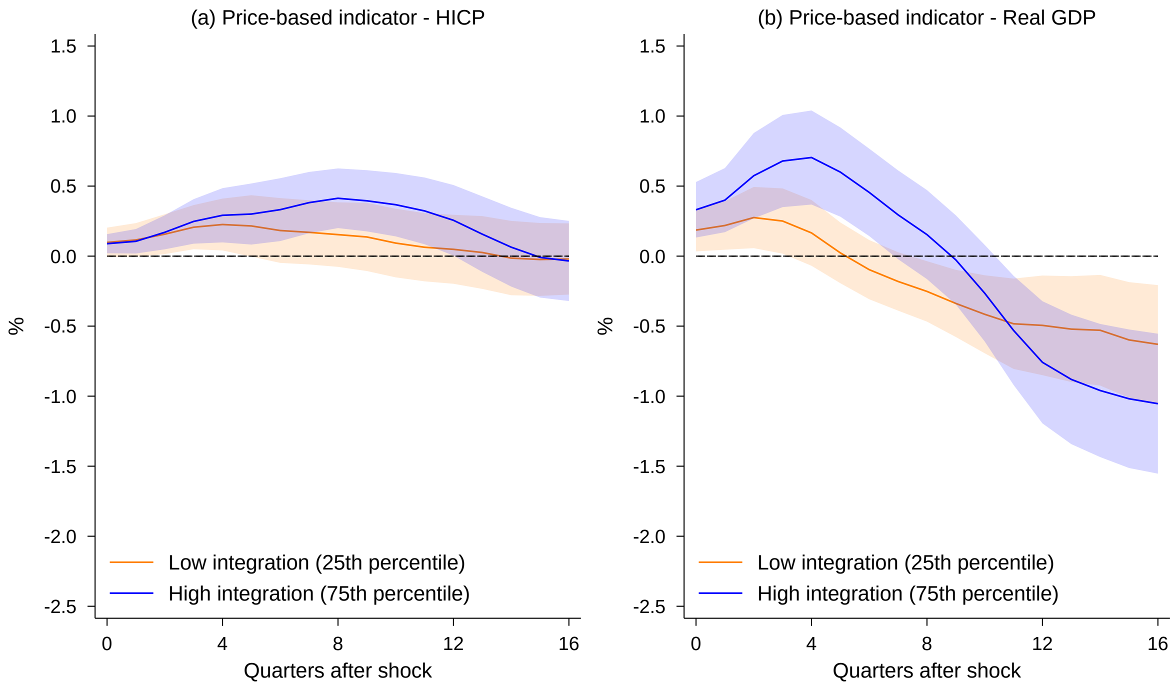

*Notes:* The two regimes of financial integration (low versus high) are defined based on a dummy created using data on the quantity-based composite indicator. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (below median, orange line) and high (above median, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

### B.3.4 Price- versus quantity-based composite indicator

We replace the quantity-based composite indicator from Section 2.1 with a price-based one, also from Hoffmann, Kremer and Zaharia (2020), to test whether the results depend on the measure of financial integration. The baseline relies on cross-border asset holdings (quantities), while the price-based indicator measures integration through cross-country convergence of asset prices. We apply the same specification from Equation 2.

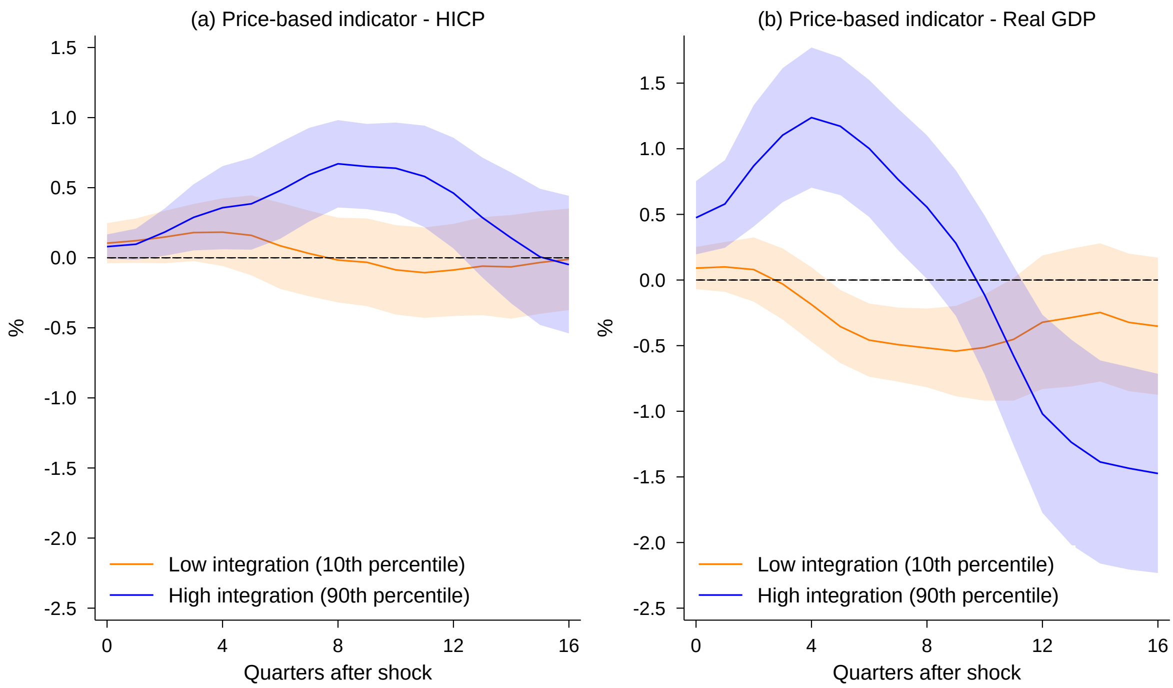
Figure B.16 shows that the amplification pattern is reproduced: monetary shocks have stronger effects on consumer prices and output under high integration. To assess robustness in the tails of the integration distribution, Figure B.17 reports responses at the 10th and 90th percentiles of the price-based indicator; the results are qualitatively unchanged.

**Figure B.16:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Price-based composite indicator from Hoffmann, Kremer and Zaharia (2020) is used instead of the quantity-based indicator. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

**Figure B.17:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration

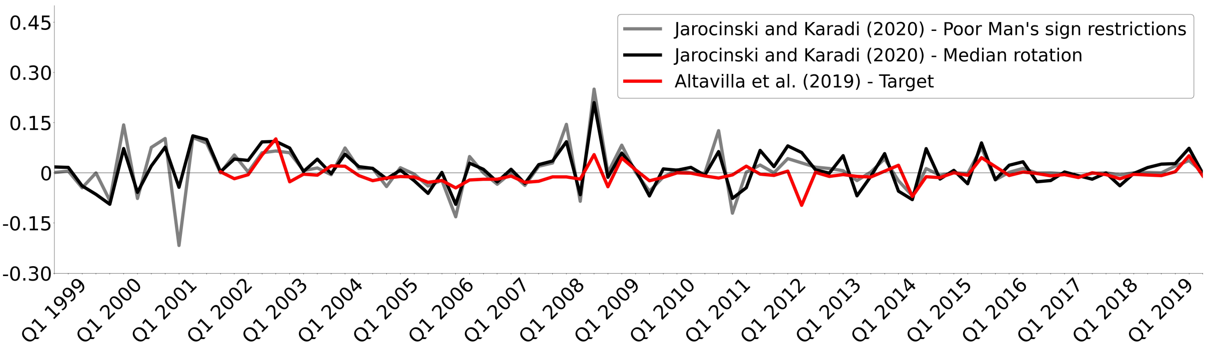



*Notes:* Price-based composite indicator from Hoffmann, Kremer and Zaharia (2020) is used instead of the quantity-based indicator, and the 10th and 90th percentile are considered (instead of 25th and 75th). Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (10th percentile, orange line) and high (90th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

### B.3.5 Alternative monetary policy shocks

Jarociński and Karadi (2020) estimate monetary policy and central bank information shocks using two methods: a simple ("Poor Man's") sign restrictions approach and a median rotation approach. The first method identifies monetary policy shocks with basic sign restrictions (e.g., higher interest rates with falling stock prices) and can also detect information shocks (where interest rates and stock prices move together), but it uses a simple high-frequency framework that may not fully separate the two. The median rotation approach better distinguishes monetary policy shocks (opposite movements of interest rates and stocks) from information shocks (same-direction movements) by choosing the median response from a set of valid rotations. Using shocks from this second method, we re-estimate the impulse responses; Figure B.19 shows that the amplification pattern holds.

**Figure B.18:** High-frequency monetary shocks from the literature



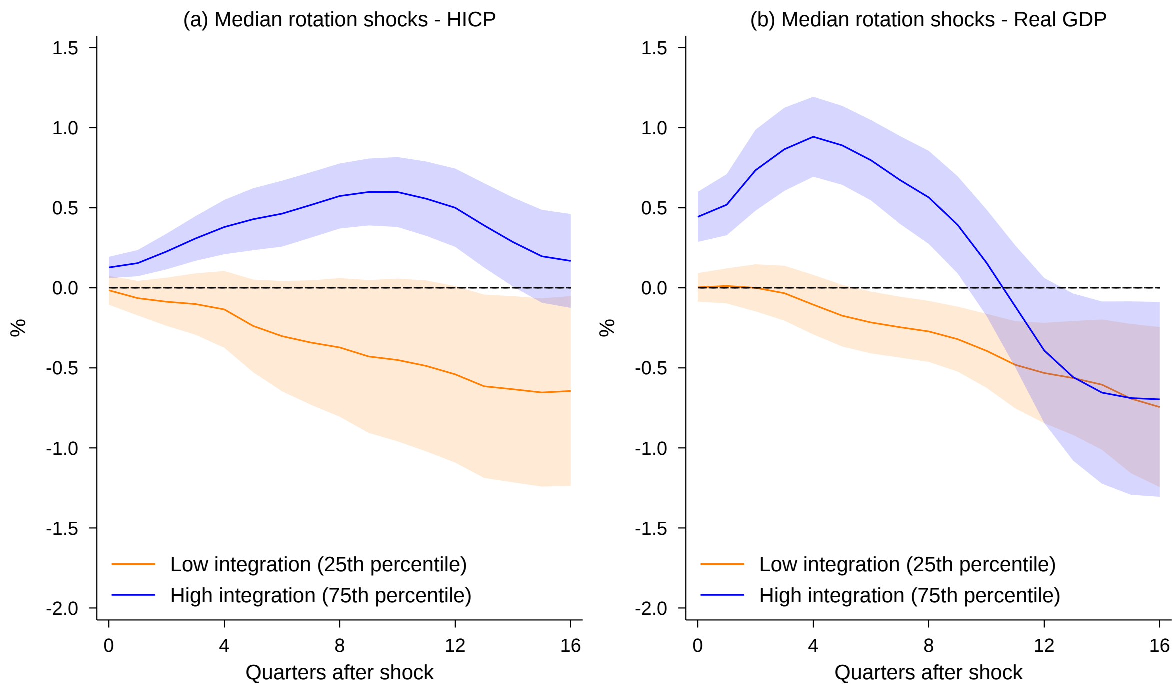

*Notes:* Monetary policy shocks from Jarociński and Karadi (2020) and target shocks from Altavilla et al. (2019).

Another approach to identifying monetary policy shocks comes from Altavilla et al. (2019). They break down ECB policy surprises into target, timing, forward guidance, and quantitative easing factors using high-frequency changes in OIS yields. Target shocks, which reflect surprises in short-term policy rates, are identified from intraday yield curve shifts around ECB announcements, especially press releases.

Figure B.18 compares the two shock measures from Jarociński and Karadi (2020) with the target shocks from Altavilla et al. (2019). The two Jarociński and Karadi (2020) shock series are highly correlated, while the Altavilla et al. (2019) target shocks have smaller amplitude but co-move directionally with the JK shocks. The smaller magnitude of the target shocks reflects the fact that they are identified from changes only in short-term interest rates, whereas the Jarociński and Karadi (2020) shocks are based on changes in Eonia interest rate swaps at short to medium maturities, which incorporate both current and expected future policy moves. The two approaches also differ in their treatment of central bank information effects, which accounts for residual discrepancies despite the broadly similar co-movement.

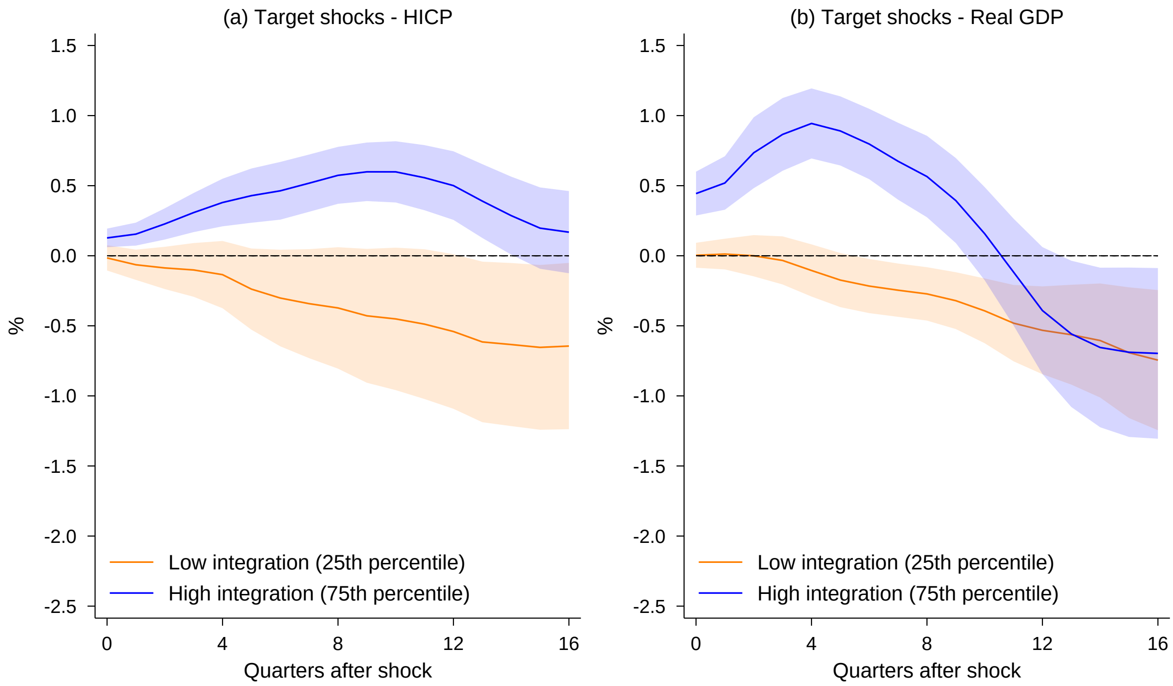
Figure B.20 reports the impulse responses using the target shocks. The amplification pattern is preserved across all three shock measures.

**Figure B.19:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* The alternative shocks from Jarociński and Karadi (2020) are used. Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

**Figure B.20:** Responses of HICP and real GDP to a monetary policy shock under low and high levels of financial integration




*Notes:* Target shocks from Altavilla et al. (2019) are used instead of shocks from Jarociński and Karadi (2020). Impulse responses in levels (percentage terms) to a one-standard-deviation expansionary monetary shock under low (25th percentile, orange line) and high (75th percentile, blue line) levels of financial integration. Shaded areas denote 68% confidence intervals computed using Driscoll-Kraay standard errors.

# C A Simple Model of Financial Integration and Monetary Transmission

This appendix develops a stylized model that formalizes the mechanism described in the main text. The amplification result itself—that stronger pass-through increases the macroeconomic response—is intuitive. The model's value lies in the precise functional form it generates: a linear interaction structure that maps directly to the empirical specification in Equation 2, a core–periphery ranking determined by intrinsic pass-through heterogeneity, and an analytical prediction for the $R^2$ of yield regressions across integration regimes (Section 4.4), conditional on the relative variances of common and idiosyncratic shocks. Each of these predictions is testable and finds support in the data.

## C.1 Environment

Consider a monetary union of two countries $j \in \{1, 2\}$ with a single central bank that sets the policy rate $i_t$. Each country has a representative household and a representative firm. We impose four assumptions on the baseline model; a fifth is introduced in the extension (Section C.6).

**A1** (*Symmetric New Keynesian block*). Both countries follow a standard log-linearized New Keynesian structure (Gali and Monacelli, 2008) with common structural parameters $(\sigma, \beta, \kappa)$. All variables denote log-deviations from steady state. Country $j$'s output gap $y_{j,t}$ and inflation $\pi_{j,t}$ satisfy (since **A1** abstracts from supply shocks, natural output is constant and $y_{j,t}$ equivalently denotes the log-deviation of output from steady state):

$$y_{j,t} = \mathbb{E}_t y_{j,t+1} - \sigma^{-1}\left(r^L_{j,t} - \mathbb{E}_t \pi_{j,t+1}\right), \tag{C.1}$$

$$\pi_{j,t} = \beta\, \mathbb{E}_t \pi_{j,t+1} + \kappa\, y_{j,t}, \tag{C.2}$$

where $\sigma > 0$ is the inverse elasticity of intertemporal substitution, $\beta \in (0, 1)$ is the discount factor, $\kappa > 0$ is the slope of the Phillips curve, and $r^L_{j,t}$ is the lending rate relevant for country $j$'s spending decisions.

The assumption that both countries share $(\sigma, \beta, \kappa)$ ensures that all cross-country heterogeneity in monetary transmission originates in the financial block described next.

## C.2 Financial Intermediation and Pass-Through

The policy rate $i_t$ is the rate at which the central bank provides funds to the banking system. Under financial fragmentation, however, this rate does not flow uniformly to all banks. Vari (2020) provides the key theoretical result: when the interbank market fragments along national lines—due to counterparty risk that differs across countries—banks in different member states face different effective funding costs. In his model, peripheral banks, perceived as riskier counterparties, must pay a premium in the interbank market or borrow directly from the central bank at the marginal lending facility rate; core banks accumulate excess liquidity from deposit inflows and fund near the deposit facility rate. The result is a two-tier interbank market in which the single policy rate produces country-specific funding costs $r_{j,t}^{\text{local}}$ that can diverge substantially from $i_t$. Abbassi et al. (2022) confirm this prediction empirically: during the Lehman and sovereign crises, borrowers from peripheral countries paid up to 100 basis points more than core borrowers for identical overnight loans on the same day. Financial integration reverses this process: as cross-border interbank flows resume, banks can access funding at rates closer to the policy rate, closing the gap between $r_{j,t}^{\text{local}}$ and $i_t$.

We capture this mechanism in reduced form:

**A2** (*Interbank segmentation and pass-through*). The lending rate in country $j$ is

$$r_{j,t}^{L} = \lambda_j(\theta)\, i_t + \left[1 - \lambda_j(\theta)\right] r_{j,t}^{\text{local}}, \tag{C.3}$$

where the pass-through coefficient is

$$\lambda_j(\theta) = \theta + (1-\theta)\,\psi_j, \qquad \psi_j \in [0,1). \tag{C.4}$$

Here $\theta \in [0,1]$ indexes the degree of union-wide financial integration. We do not derive a closed-form mapping from Vari's (2020) structural parameters; rather, $\theta$ is a reduced-form summary of the forces—counterparty risk, cross-border interbank activity—that Vari (2020) models structurally, chosen for analytical tractability. Qualitatively, higher $\theta$ corresponds to a smaller counterparty risk wedge in his framework. The parameter $\psi_j$ captures country $j$'s intrinsic pass-through under financial autarky ($\theta = 0$), determined by the severity of the counterparty risk premium its banks face in the interbank market. The strict inequality $\psi_j < 1$ excludes perfect pass-through under complete segmentation, which would eliminate any role for cross-border arbitrage.

Two polar cases clarify the role of $\theta$. When $\theta = 1$, interbank markets are fully integrated: cross-border lending eliminates the funding cost wedge and $\lambda_j = 1$ for all $j$, so the policy rate passes through completely. When $\theta = 0$, interbank markets are fully segmented: banks fund locally at $r_{j,t}^{\text{local}}$ and $\lambda_j = \psi_j < 1$, so pass-through is incomplete and heterogeneous across countries.

The implied spread between the lending rate and the policy rate follows directly from (C.3):

$$r_{j,t}^{L} - i_t = \left[1 - \lambda_j(\theta)\right]\left(r_{j,t}^{\text{local}} - i_t\right). \tag{C.5}$$

This is a multi-country extension of the intermediation wedge identified by Gertler and Karadi (2011) and Curdia and Woodford (2010), who show that frictions in financial intermediation—balance sheet constraints and costly intermediation, respectively—attenuate the pass-through of monetary policy to the real economy. In our setting, the wedge arises from interbank market segmentation (Vari, 2020): banks fund at $r_{j,t}^{\text{local}} \neq i_t$ because cross-border interbank lending is impaired by counterparty risk. Financial integration compresses the wedge by restoring cross-border flows.

## C.3 Main Results

Define the average output and inflation responses as $\bar{y}_t = \frac{1}{2}(y_{1,t} + y_{2,t})$ and $\bar{\pi}_t = \frac{1}{2}(\pi_{1,t} + \pi_{2,t})$, corresponding to what the panel specification in Equation 2 estimates with country fixed effects. To isolate the pass-through channel, we impose:

**A3** (*Partial equilibrium*). The monetary policy shock $\Delta i_t$ is unanticipated. On impact, expectations and local funding rates do not respond: $\Delta \mathbb{E}_t y_{j,t+1} = \Delta \mathbb{E}_t \pi_{j,t+1} = \Delta r^{\text{local}}_{j,t} = 0$.

This is a comparative-static simplification, not an equilibrium restriction: it isolates the pass-through scaling by ensuring that the impact response depends only on the lending rate channel $\lambda_j(\theta)$, not on the general-equilibrium adjustment of expectations. In a full solution of the NK model, a persistent monetary shock would move expectations contemporaneously; we reintroduce these dynamics in Section C.4.

**Proposition 1** (Amplification)**.** *Under **A1–A3**, the impact responses of average output and average inflation to a monetary policy shock $\Delta i_t$ are*

$$\Delta \bar{y}_t = -\frac{1}{\sigma}\,\bar{\lambda}(\theta)\,\Delta i_t, \tag{C.6}$$

$$\Delta \bar{\pi}_t = -\frac{\kappa}{\sigma}\,\bar{\lambda}(\theta)\,\Delta i_t, \tag{C.7}$$

*where $\bar{\lambda}(\theta) = \theta + (1-\theta)\bar{\psi}$ and $\bar{\psi} = \frac{1}{2}(\psi_1 + \psi_2)$. Both responses are strictly increasing in financial integration:*

$$\frac{\partial |\Delta \bar{y}_t|}{\partial \theta} = \frac{1}{\sigma}(1-\bar{\psi})\,|\Delta i_t| > 0, \qquad \frac{\partial |\Delta \bar{\pi}_t|}{\partial \theta} = \frac{\kappa}{\sigma}(1-\bar{\psi})\,|\Delta i_t| > 0. \tag{C.8}$$

The joint amplification of output and prices is not a coincidence—it is a structural feature of the New Keynesian framework. The Phillips curve ties inflation to the output gap, so any friction that attenuates the output response to monetary policy necessarily attenuates the inflation response by the same channel. Incomplete pass-through under financial fragmentation is precisely such a friction: it weakens the real rate movement that drives the output gap, which in turn dampens the inflationary pressure that the Phillips curve translates from the output gap. This is consistent with the empirical finding that financial integration amplifies the transmission of monetary policy to both HICP and real GDP (Section 4).

*Proof. Step 1 (Country-level output response).* From the IS curve (C.1), applying **A3** ($\Delta \mathbb{E}_t y_{j,t+1} = \Delta \mathbb{E}_t \pi_{j,t+1} = 0$):

$$\Delta y_{j,t} = -\sigma^{-1}\,\Delta r^L_{j,t}.$$

*Step 2 (Country-level lending rate response).* From the lending rate equation (C.3), applying **A3** ($\Delta r^{\text{local}}_{j,t} = 0$):

$$\Delta r^L_{j,t} = \lambda_j(\theta)\,\Delta i_t.$$

*Step 3 (Averaging).* Substituting Step 2 into Step 1: $\Delta y_{j,t} = -\sigma^{-1}\lambda_j(\theta)\,\Delta i_t$. Averaging over $j \in \{1,2\}$:

$$\Delta \bar{y}_t = -\sigma^{-1}\,\bar{\lambda}(\theta)\,\Delta i_t,$$

where $\bar{\lambda}(\theta) = \frac{1}{2}[\lambda_1(\theta) + \lambda_2(\theta)] = \frac{1}{2}[(\theta + (1-\theta)\psi_1) + (\theta + (1-\theta)\psi_2)] = \theta + (1-\theta)\bar{\psi}$.

*Step 4 (Amplification).* Since $\theta \in [0,1]$ and $\psi_j \in [0,1)$, we have $\bar{\lambda}(\theta) \geqslant 0$, so $|\Delta \bar{y}_t| = \sigma^{-1}\bar{\lambda}(\theta)|\Delta i_t|$. Differentiating with respect to $\theta$:

$$\frac{\partial |\Delta \bar{y}_t|}{\partial \theta} = \sigma^{-1}\frac{\partial \bar{\lambda}}{\partial \theta}|\Delta i_t| = \sigma^{-1}(1-\bar{\psi})|\Delta i_t|.$$

Since $\psi_j < 1$ for each $j$, $\bar{\psi} < 1$, so $1 - \bar{\psi} > 0$.

*Step 5 (Inflation).* From the Phillips curve (C.2), applying **A3** ($\Delta \mathbb{E}_t \pi_{j,t+1} = 0$): $\Delta \pi_{j,t} = \kappa \Delta y_{j,t}$. Averaging: $\Delta \bar{\pi}_t = \kappa \Delta \bar{y}_t = -(\kappa/\sigma)\bar{\lambda}(\theta)\Delta i_t$. The amplification follows by the same differentiation as Step 4, scaled by $\kappa$. ■

**Corollary 1** (Core–Periphery Heterogeneity)**.** *Under* ***A1*–*A3****, if $\psi_{core} > \psi_{periphery}$, then the amplification of both output and inflation in the periphery is strictly larger than in the core:*

$$\frac{\partial |\Delta y_{periphery,t}|}{\partial \theta} > \frac{\partial |\Delta y_{core,t}|}{\partial \theta}, \qquad \frac{\partial |\Delta \pi_{periphery,t}|}{\partial \theta} > \frac{\partial |\Delta \pi_{core,t}|}{\partial \theta}. \tag{C.9}$$

*Proof.* From Steps 1–2 of the proof of Proposition 1, the country-level output response is $\Delta y_{j,t} = -\sigma^{-1}\lambda_j(\theta)\Delta i_t$, and from Step 5, $\Delta \pi_{j,t} = \kappa \Delta y_{j,t}$. Since $\lambda_j(\theta) \geqslant 0$, the output sensitivity is:

$$\frac{\partial |\Delta y_{j,t}|}{\partial \theta} = \sigma^{-1}(1-\psi_j)|\Delta i_t|,$$

which is strictly decreasing in $\psi_j$. The inflation sensitivity is $\kappa$ times the output sensitivity, so the ranking is preserved. Since $\psi_{core} > \psi_{periphery}$, both inequalities follow. ■

The sensitivity $\partial|\Delta y_{j,t}|/\partial\theta = \sigma^{-1}(1-\psi_j)|\Delta i_t|$ is decreasing in $\psi_j$: countries whose banks face a larger counterparty risk premium (lower $\psi_j$) have a larger intermediation wedge under segmentation, so integration changes their monetary transmission more. This differential applies symmetrically: a periphery country experiences a larger contraction from a tightening shock and a larger expansion from an easing shock. The model does not derive the ranking $\psi_{core} > \psi_{periphery}$ from primitives; rather, it shows that *if* intrinsic pass-through differs across country groups, *then* integration amplifies transmission differentially. The direction of this heterogeneity is an empirical input, supported by the micro-evidence on more efficient banking intermediation in core member states documented by Hristov, Hülsewig and Wollmershäuser (2014) and Von Borstel, Eickmeier and Krippner (2016), and consistent with the subsample estimates in Section 4.3.

## C.4 Dynamic Extension and Mapping to the Empirical Specification

The results above characterize the impact response under **A3**. At longer horizons, the New Keynesian dynamics in (C.1)–(C.2) propagate the initial impulse through the general-equilibrium response of expectations, output, and inflation. **A3** fixes the local rate on impact only; at longer horizons, the local rate may respond to domestic conditions, requiring the additional restriction in **A4**. To extend the multiplicative separability of the impact result to the full dynamic path, we impose:

**A4** (*Exogenous local rates*). The local funding rate $r_{j,t}^{\text{local}}$ does not respond to the monetary policy shock at any horizon. This ensures that $\theta$ enters the system exclusively through the initial scaling of the lending rate in (C.3) and does not affect the propagation dynamics.

Under **A1–A2** and **A4**, the log-linearized system is linear and $\theta$ enters only through (C.3), scaling the impulse to the real lending rate without affecting the propagation coefficients in (C.1)–(C.2). Let $f_h^y$ and $f_h^\pi$ denote the dynamic multipliers at horizon $h$ for output and inflation, respectively—the impulse response functions of the average economy to a unit impulse in the real lending rate. At $h = 0$, the partial-equilibrium logic of **A3** gives $f_0^y = -1/\sigma$ (the contemporaneous IS elasticity, holding expectations fixed). At $h > 0$, the $f_h$ depend on $(\sigma, \beta, \kappa)$ and on the persistence of the shock through the full rational-expectations solution of (C.1)–(C.2), and satisfy $f_h^y, f_h^\pi < 0$ for all $h \geqslant 0$ under a contractionary shock in standard parameterizations. The impulse response functions at horizon $h$ are:

$$\text{IRF}_h^y = f_h^y \cdot \bar{\lambda}(\theta) \cdot \Delta i_t, \qquad \text{IRF}_h^\pi = f_h^\pi \cdot \bar{\lambda}(\theta) \cdot \Delta i_t. \tag{C.10}$$

The dynamic profiles $f_h^y$ and $f_h^\pi$ differ—inflation is forward-looking in the output gap, so its dynamic path need not be a simple scaling of the output path—but $\theta$ enters multiplicatively in both cases, so the amplification effect scales the entire dynamic response path for both variables. Expanding $\bar{\lambda}(\theta) = \bar{\psi} + (1 - \bar{\psi})\,\theta$ (using output for concreteness; the inflation mapping is identical with $f_h^\pi$ replacing $f_h^y$):

$$\text{IRF}_h = \underbrace{f_h\,\bar{\psi}}_{\equiv\ \beta_h} \cdot \Delta i_t \;+\; \underbrace{f_h\,(1 - \bar{\psi})}_{\equiv\ \delta_h} \cdot \left(\Delta i_t \times \theta\right), \tag{C.11}$$

where $\beta_h$ denotes the regression intercept from the empirical specification (not the discount factor in (C.2)), which takes exactly the form of Equation 2—up to fixed effects and controls—with $\theta$ playing the role of $FI_{t-1}$. The model delivers three predictions for the estimated coefficients:

1. $\delta_h$ and $\beta_h$ have the same sign (amplification, not sign reversal), since $\delta_h = \beta_h \cdot (1 - \bar{\psi})/\bar{\psi}$ for $\bar{\psi} > 0$;

2. the amplification is proportional to $(1 - \bar{\psi})$, the average pass-through gap under autarky; and

3. the ratio $\delta_h/\beta_h = (1 - \bar{\psi})/\bar{\psi}$ is constant across horizons $h$ under **A4**—this is the model's sharpest testable prediction. If the ratio varies across horizons, this would indicate that Assumption 4 is violated—i.e., that the local rate response is not horizon-invariant—rather than contradicting the core amplification mechanism. In the data, $\hat{\beta}_h$ (the effect at $FI = 0$, outside the data support) is imprecisely estimated—never significant at the 90% confidence level for HICP—making the ratio unreliable at many horizons. For real GDP, where both coefficients are precisely estimated at short-to-medium horizons, the ratio is stable at $h = 1$ and $h = 4$ ($-4.75$ and $-4.70$, respectively) but drifts toward zero at longer horizons ($-3.71$ at $h = 8$). The negative sign of these ratios reflects the empirical sign convention: $\hat{\beta}_h$ captures the direct effect at $FI = 0$, which is small

and occasionally positive because integration is never truly zero in the data, while $\hat{\delta}_h$ is negative (amplifying the expansionary effect of an easing shock). The model predicts a positive ratio under the assumption that $\beta_h < 0$ (a contractionary shock reduces output even at $FI = 0$), which maps to the empirical estimates only after accounting for the sign-flip induced by the lack of data support at zero integration. The key testable content of Prediction 3 is therefore the approximate constancy of the ratio across horizons, which the data support at short-to-medium horizons. The drift at longer horizons is consistent with a gradual violation of **A4** as local funding rates begin responding to the shock.

### C.5 Mapping to the Mechanism Evidence

The model also rationalizes the evidence on interest rate pass-through uniformity presented in Section 4.4. Consider regressing country-level yield changes $\Delta r^L_{j,t}$ on the policy rate change $\Delta i_t$. From (C.3):

$$\Delta r^L_{j,t} = \lambda_j(\theta)\,\Delta i_t + \left[1 - \lambda_j(\theta)\right]\Delta r^{\text{local}}_{j,t}.$$

If $\Delta i_t$ and $\Delta r^{\text{local}}_{j,t}$ are uncorrelated, the $R^2$ of this regression is (the zero-covariance assumption is a simplification; during the sovereign debt crisis, local rates moved against the policy rate, which would widen the $R^2$ gap beyond what the formula predicts; in tranquil periods, partial positive covariance would compress the gap; the prediction is therefore qualitative):

$$R^2 = \frac{\lambda_j(\theta)^2\,\mathrm{Var}(\Delta i_t)}{\lambda_j(\theta)^2\,\mathrm{Var}(\Delta i_t) + \left[1 - \lambda_j(\theta)\right]^2 \mathrm{Var}(\Delta r^{\text{local}}_{j,t})}. \tag{C.12}$$

When $\theta \to 1$, $\lambda_j \to 1$ and the local rate term vanishes, so $R^2 \to 1$: all variation in lending rates is driven by the common policy rate. When $\theta \to 0$, $\lambda_j = \psi_j < 1$ and local rate variation reduces $R^2$ in proportion to $\mathrm{Var}(\Delta r^{\text{local}}_{j,t})/\mathrm{Var}(\Delta i_t)$. The prediction is qualitative—the exact $R^2$ depends on the relative variances of policy and local shocks—but it is consistent with the evidence in Section 4.4, where the $R^2$ of yield changes on the policy rate is 84% under high financial integration and 52% under low integration. Two features distinguish the model formula from the empirical exercise. First, the formula characterizes a single country, whereas the empirical $R^2$ pools across countries with heterogeneous $\lambda_j$ and local rate variances. Second, the formula applies to the lending rate $r^L_{j,t}$, whereas the empirical $R^2$ values of 84% and 52% are from regressions of *sovereign* yield changes on the OIS rate. Neither gap affects the qualitative prediction. Pooling preserves the comparative static with respect to $\theta$, and sovereign yields reflect the same forces—cross-border arbitrage, counterparty risk premia, and market segmentation—that govern $\lambda_j(\theta)$ in the model. The prediction that higher $\theta$ raises the share of yield variation explained by the common shock holds for any financial variable whose pass-through from the policy rate is governed by integration.

## C.6 Extension: Credit and Portfolio Channels

The baseline model captures a single transmission channel: the interest rate pass-through from the policy rate to national lending rates. The empirical channel decomposition in Section 4.4.2, however, shows that banking market integration (a credit supply channel) and equity market integration (a portfolio channel) contribute independently to the amplification of monetary transmission. We now show that the model's structure generalizes naturally to accommodate these additional channels.

**A5** (*Direct financial channels*). Beyond the interest rate channel, the monetary policy shock $\Delta i_t$ affects country $j$'s output through additional financial channels—cross-border credit supply and cross-border portfolio rebalancing—whose strength depends on financial integration. These channels are captured by augmenting the IS curve with a reduced-form impact-response term (this should be understood as a contemporaneous response specification rather than a structural Euler equation modification):

$$y_{j,t} = \mathbb{E}_t y_{j,t+1} - \sigma^{-1}\left(r^L_{j,t} - \mathbb{E}_t \pi_{j,t+1}\right) - \phi_j(\theta)\,\Delta i_t, \tag{C.13}$$

where $\phi_j(\theta) \geqslant 0$ captures the combined sensitivity of credit supply and portfolio valuation to the policy rate, and

$$\phi_j(\theta) = \theta\,\bar{\phi} + (1-\theta)\,\phi_j^0, \qquad \bar{\phi} > \phi_j^0 \geqslant 0. \tag{C.14}$$

Under full integration ($\theta = 1$), all countries experience the same direct financial stimulus $\bar{\phi}\,\Delta i_t$; under segmentation ($\theta = 0$), the stimulus is $\phi_j^0\,\Delta i_t$, which is weaker and heterogeneous across countries.

The two components of $\phi_j$ correspond to distinct economic mechanisms:

- *Credit supply channel* (banking integration): When banking markets are integrated, a rate cut induces cross-border credit expansion as banks reallocate lending across member states through internal capital markets (Peek and Rosengren, 1997; Cetorelli and Goldberg, 2012). Under segmentation, peripheral banks face tighter balance sheet constraints and cannot expand lending as freely, weakening the credit supply response (Bruno and Shin, 2015; Kalemli-Ozcan, Papaioannou and Peydró, 2013).
- *Portfolio rebalancing channel* (equity integration): When equity markets are integrated, a rate cut produces uniform asset valuation gains across countries, generating wealth effects on consumption and reinforcing the demand stimulus (Bernanke and Kuttner, 2005; Fratzscher, Lo Duca and Straub, 2016). Under segmentation, valuation gains remain localized, attenuating the cross-border transmission.

Under **A1–A5**, the impact response of average output generalizes Proposition 1:

$$\Delta\bar{y}_t = -\left[\sigma^{-1}\,\bar{\lambda}(\theta) + \bar{\phi}(\theta)\right]\Delta i_t, \tag{C.15}$$

where $\bar{\phi}(\theta) = \theta\,\bar{\phi} + (1-\theta)\,\bar{\phi}^0$ and $\bar{\phi}^0 = \frac{1}{2}(\phi_1^0 + \phi_2^0)$. The total amplification with respect to integration is

$$\frac{\partial|\Delta\bar{y}_t|}{\partial\theta} = \Big[\ \underbrace{\sigma^{-1}(1-\bar{\psi})}_{\text{rate pass-through}} \ + \ \underbrace{(\bar{\phi}-\bar{\phi}^0)}_{\text{credit + portfolio}}\ \Big]\,|\Delta i_t| > 0. \tag{C.16}$$

The first term is the interest rate channel from Proposition 1; the second term captures the additional amplification from the credit and portfolio channels. The linear interaction structure of Equation 2 is preserved: $\theta$ still enters multiplicatively with $\Delta i_t$, and the mapping to the empirical specification in (C.11) continues to hold with $\delta_h$ now reflecting the combined contribution of all three channels.

This decomposition maps to the empirical findings in Section 4.4.2: the composite financial integration indicator captures $\theta$ as a common driver of all three channels, while the sub-index decomposition isolates the relative contribution of each market segment. The finding that banking integration dominates for GDP and equity integration dominates for HICP is consistent with the credit channel operating primarily through investment and the portfolio channel through consumption—channels that have different relative impacts on output versus prices along the Phillips curve.

### C.7 Discussion

The model formalizes the verbal argument in the main text using minimal structure: the standard New Keynesian macro block (**A1**) combined with a reduced-form representation of interbank market segmentation (**A2**) and, in the extension, direct financial channels (**A5**). Despite its simplicity, it delivers all key empirical predictions of the paper—amplification, core–periphery heterogeneity, the pass-through mechanism, and the multi-channel structure identified in the sub-index decomposition.

The model builds directly on Vari (2020), who shows that interbank market fragmentation—driven by counterparty risk that differs across countries—causes the policy rate to transmit heterogeneously to bank funding costs. Our contribution is to trace this mechanism through to its macroeconomic consequences: by embedding Vari's (2020) interbank segmentation in a standard New Keynesian macro block, we show that it attenuates the aggregate response of output and prices to monetary policy, derive the specific functional form of the amplification (linear in $\theta$, proportional to the average pass-through gap $1-\bar{\psi}$), and map it to a panel local projections specification with an interaction term. Vari (2020) stops at the interbank market—showing that rates diverge; we show what this divergence implies for aggregate monetary transmission, and how to test it. This also connects to the intermediation wedge that Gertler and Karadi (2011) and Curdia and Woodford (2010) identified as a key friction in monetary transmission: our model shows that this wedge arises from the assumed interbank segmentation structure, following Vari (2020), in a monetary union. The model is also related to Wu, Xie and Zhang (2024) and Bianchi and Coulibaly (2024), who show that financial integration amplifies monetary transmission in settings with independent monetary policies and flexible exchange rates. Our contribution is to isolate the within-union channel: a common policy rate that is transmitted heterogeneously across countries depending on the degree of interbank market integration, without requiring exchange rate or terms-of-trade adjustments.

We note four limitations. First, the model treats $\theta$ as predetermined with respect to the monetary policy shock. This is by design: it mirrors the empirical specification in Equation 2, where financial integration enters as $FI_{t-1}$, lagged one quarter relative to the shock. The model therefore characterizes transmission *conditional on* the prevailing level of integration, without requiring that integration itself be exogenous to all macroeconomic forces. Second, **A4** holds local funding rates fixed; in practice, local rates may respond to the monetary policy shock. During episodes where local rates move *against* the policy rate—as occurred during the sovereign debt crisis, when ECB easing coincided with widening periphery spreads—the countervailing channel would generate even stronger attenuation under low integration than the model predicts, making our estimates a lower bound for those periods. In tranquil periods where local rates partially co-move with the policy rate, the bound does not apply and the model may overstate the amplification difference. Third, the extension in Section C.6 treats the credit and portfolio channels in reduced form; a structural model of cross-border banking with endogenous balance sheet constraints (Gertler and Karadi, 2011; Bruno and Shin, 2015) and cross-border portfolio choice (Coeurdacier and Rey, 2013) would provide micro-foundations for $\phi_j(\theta)$ and potentially generate richer predictions about which channel dominates under different macroeconomic conditions. Fourth, the framework abstracts from the zero lower bound and unconventional policy instruments. Each of these extensions would enrich the model at the cost of analytical tractability, and we leave them for future work.