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The effect of fragmentation risk on monetary conditions in the euro area

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ABSTRACT

This paper measures the output effects of financial fragmentation in the euro area by estimating an extended IS curve. Using a panel approach, we find that two fragmentation measures are significantly related to the output gap: sovereign spreads and spreads in the long-term cost of borrowing of the private sector. We use these output effects to construct a Monetary Conditions Index (MCI) for euro area countries. This index summarizes the combined effect of the monetary policy stance and financial fragmentation. We show that the MCI approach is well-suited to capture cross-country differences in a fragmentation-enhanced measure of the monetary policy stance. Using this metric, we find that during the sovereign debt crisis, the cross-country dispersion of MCI's based on sovereign spreads was much larger than that based on the private cost of borrowing. We also show that convergence is slower for MCI's based on sovereign spreads. We conclude that the causes of fragmentation in monetary conditions may change over time, and that this has implications for the appropriate policy response.

1. Introduction

In 2022, the European Central Bank (ECB) announced the introduction of an anti-fragmentation tool (ECB 2022a). This new monetary policy instrument, labelled "Transmission Protection Instrument" (TPI), should enable the ECB to control sovereign spreads of euro area (EA) countries, through the purchase of government bonds from countries whose interest rates are deemed to be out of step with their macroeconomic fundamentals. In the way, risk premia arising from unwarranted negative sentiment in EA bond markets could be reduced. The ECB's justification for introducing the TPI is that diverging yields on sovereign debt may hamper the transmission of monetary policy across the EA and increase the risk of financial fragmentation. If, for example, monetary policy tightening would result in private lending rates rising more sharply in highly indebted EA countries, monetary policy would no longer have the same effect across the union. This would compromise the singleness of monetary policy and be a concern for the ECB.

The introduction of the TPI has been met with criticism from within the economics discipline, see e.g. Bernoth et al. (2022a), Buiter (2022), Feld et al. (2022) and Marsh (2022). Arguments against the TPI are that it would shield EA countries from market discipline, thus undermining the incentives for sound national fiscal policies, that the appropriate size of risk premia is hard to establish and that the TPI seems to be more an instrument of fiscal instead of monetary policy. More generally, one can question whether the ECB should take ownership of the problem of EA fragmentation. According to Wyplosz (2023):

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> "They (i.e. central banks) have acquired the status of problem-solvers of first resort, but many of the current challenges lay outside their competence. This is the case of the various heterogeneities that have emerged across and within euro area member countries. The ECB has no instrument and no legal basis to deal with these heterogeneities". (p. 21)

In anticipation of some of these concerns, the ECB has decided that two requirements need to be met before the TPI can be activated. First, TPI activation only takes place "to counter unwarranted, disorderly market dynamics that pose a serious threat to the transmission of monetary policy across the euro area" (ECB 2022b). This implies that the ECB will purchase assets "in jurisdictions experiencing a deterioration in financing conditions not warranted by country-specific fundamentals" (ECB 2022b). Second, the ECB will assess whether "jurisdictions in which the Eurosystem may conduct purchases under the TPI pursue sound and sustainable fiscal and macroeconomic policies" (ECB 2022b). For this assessment, the ECB considers four criteria: 1) compliance with the EU fiscal framework; 2) absence of severe macroeconomic imbalances; 3) fiscal sustainability; and 4) sound and sustainable macroeconomic policies.

While the TPI has not yet been activated, the debate on whether spread control should be a task of the ECB will not go away. The question therefore remains whether the issue of fragmentation risk warrants the introduction of a new dedicated monetary policy instrument. This paper aims to contribute to this debate by exploring the output effects of fragmentation and by proposing a new metric.

From a macroeconomic perspective, the main risk of financial fragmentation is that diverging interest rates set in motion divergences in output and inflation across EA countries. In a New Keynesian model, this works via the "Investment-Savings" (IS) curve, which models how monetary policy affects the output gap. The effect on inflation then follows from the Phillips curve. Any attempt to quantify fragmentation risk should therefore not only take into account the divergence in interest rates itself, but also its effect on output. The stronger the effect of financial fragmentation on output, the stronger the case for policy interventions aimed at reducing fragmentation. As a first contribution, this paper aims to measure the output effects of fragmentation by estimating an extended IS curve along the lines of Goodhart and Hofmann (2005a) and Hafer and Jones (2008). We do this for a panel of EA countries and for various measures of financial fragmentation.

As a second contribution, this paper proposes to summarize the combined effect of the monetary policy stance and financial fragmentation into a Monetary Conditions Index (MCI). Traditionally, a MCI has been constructed as a weighted average of interest rate and exchange rate changes. Yet the EA has become a closed but fragmented monetary union. Economic uncertainty that in the pre-euro era caused volatility in foreign exchange markets, now expresses itself in the bond markets. We therefore propose to replace the exchange rate with a measure of fragmentation risk in the MCI. An advantage of using a MCI is that the monetary policy stance and fragmentation are weighted by their effects on output. In this paper, the weights are based on the estimation of the extended IS curve. We construct MCI's for individual EA countries and examine variations in the cross-country dispersion of monetary conditions over time. We also examine whether monetary conditions in the EA converge.

As a robustness check, we examine whether the output effect of fragmentation depends on the choice of method. To this end, we estimate a vector autoregression (VAR) and calculate the impulse response of output to fragmentation risk.

Our findings are as follows. First, our estimates of the extended IS curve show that the expected real short-term interest rate has a significant negative effect on the output gap. In addition, two measures of financial fragmentation are significantly related to the output gap: sovereign spreads and spreads in the long-term cost of borrowing of the private sector. This poses a risk to the macroeconomic stability of EA countries with high spreads and corroborates the ECB's view that financial fragmentation may hinder a uniform transmission of monetary policy to the real economy. Second, we show that the MCI approach can capture cross-country differences in a fragmentation-enhanced measure of the monetary policy stance. Using this metric, we show that during the sovereign debt crisis, the cross-country dispersion of MCI's based on sovereign spreads was much larger than those based on spreads in the private cost of borrowing. We also show that convergence is slower for MCI's based on sovereign spreads.

This paper is organized as follows. Section 2 discusses the concept of fragmentation risk and illustrates the risk that fragmentation poses to the macroeconomic stability of members of a monetary union using a simple macroeconomic model. In section 3, we discuss our data and empirical approach. Section 4 reports our results. We close with concluding remarks.

2. Theoretical framework

2.1. What is fragmentation risk?

A uniform, formal definition of fragmentation risk is lacking. Most authors define fragmentation as bond market fragmentation, i.e. the divergence of interest rates on EA countries' sovereign debt (see e.g. Claeys et al. 2022; Bernoth et al. 2022b; Angeloni and Gros 2022). This definition is, however, rather narrow. For a more comprehensive understanding of the effect of fragmentation on the real economy, one would also need to look at the divergence in the borrowing costs of firms and households. Such a broader perspective would also be appropriate for the ECB, which conducts monetary policy to change lending conditions in the private sector. Empirical studies typically use nominal interest rates to measure financial fragmentation (Ceci and Pericoli 2022; Kakes and van den End 2023; Angeloni and Gros 2022). From a macroeconomic perspective, one could argue that the expected real interest rate, rather than the nominal rate, is the relevant rate for investment and savings decisions by firms and households.

A still different approach, adopted by e.g. Baele et al. (2004), De Santis (2018) and Kakes and van den End (2023), defines market fragmentation as non-fundamental fragmentation. This is the part of interest rate differentials that cannot be explained by fundamental variables. For three reasons, the current paper will not follow this approach. First, identifying the non-fundamental part of fragmentation risk is a non-trivial issue. Empirical studies typically try to establish a relationship between sovereign spreads and macroeconomic variables (Bernoth et al. 2012; Bernoth et al. 2022b). The unexplained portion of this relationship may then be construed as non-fundamental fragmentation. A methodological critique of this approach is that it involves comparing a single realization of a country's macroeconomic trajectory to market prices that, at each point in time, consider a probability distribution of all potential paths the country could have taken.[1] Instances of ex-post seemingly unjustified market sentiment may therefore be linked to fundamental uncertainty about a country's economic trajectory. For example, examining German-Italian spreads, Cadamuro and Papadia (2022) show that the spread was largest when markets worried most about Italian policies. Second, while the ECB has stated that the distinction between fundamental and non-fundamental fragmentation plays a key role in the decision to activate the TPI, it remains to be seen how this intention will hold up in practice. Monetary policymaking takes place in real time, with little time and imperfect current data for empirical analysis. Third, as we will see in section 2.2, the risk that fragmentation destabilizes the monetary union does not depend on the distinction between fundamental and non-fundamental fragmentation.

Bond market fragmentation is inherent in Europe's currency union. The EA operates as an incomplete currency union, sharing a common currency but lacking a unified fiscal policy (De Grauwe 2022). In principle, each EA country is responsible for its own national debt. The no-bailout clause and the prohibition on monetary financing have been included in the Maastricht Treaty to safeguard this principle. Consequently, investors bear credit risk on sovereign debt. Risk premia on sovereign debt may thus arise, either justified by macroeconomic fundamentals or stemming from unwarranted disorderly market dynamics. Even absent bond market fragmentation, some financial fragmentation will still occur in the EA. European markets for financial services are not fully integrated, resulting in persistent cross-country differences in the interest rates that banks pay on savings or charge for credit.

Bond market fragmentation may spill over into fragmentation in private sector borrowing costs (Wyplosz 2023). There are a number of channels through which this can happen. In principle, a higher risk premium on the sovereign debt of a EA country should not automatically translate into a higher lending rate for creditworthy firms in that country. In practice, however, spillovers may occur either via the fiscal policy response or via the financial system. When a EA country tries to maintain the confidence of the bond markets through a policy of austerity, this may negatively affect economic growth and increase credit risk in the private sector. The interconnectedness between banks and sovereigns in the EA, as evidenced by banks' exposure to domestic sovereign debt, implies that any increase in the sovereign risk premium will weaken domestic banks and increase their funding costs. As banks pass on these higher costs to their customers, private lending rates will increase. The empirical evidence suggests that, while there is a significant relationship between sovereign bond spreads and private sector borrowing costs, the relationship is not one-to-one (Arnold and van Ewijk 2014; Theobald and Tober 2020).

As bond market fragmentation is an incomplete measure of the fragmentation in lending conditions across the EA, the empirical part of this paper will use both government bond yields and private sector cost of borrowing rates to measure financial fragmentation. In addition, we will use both nominal and real interest rates.

2.2. Fragmentation risk in a small macro-model

A three-equation macroeconomic model, consisting of an *IS* curve, an expectations-augmented Phillips curve and a monetary policy (*MP*) rule, has become an established tool to explain short-run business cycle fluctuations. This paper adapts the three-equation model to analyze a single country *i* within a monetary union. We will use a small-country assumption, which implies that interest rate setting by the ECB does not react to idiosyncratic developments in country *i*. In the literature, three-equation models differ with respect to the lag structure, the expectations formation and, especially, the specification of the *MP* rule. The current model is deliberately kept simple; see Corsetti et al. (2014) for a more formal treatment.

The *IS* curve in equation (1) relates the output gap (denoted $y_{i,t}$) to the expected real interest rate ($r_{i,t}$) and a demand shock ($\epsilon_{1,i,t}$):

$$y_{i,t} = \alpha_0 - \alpha_1 r_{i,t} + \epsilon_{1,i,t}, \qquad \alpha_1 > 0 \tag{1}$$

Equation (2) is a short-run Phillips curve. It relates inflation ($\pi_{i,t}$) to expected inflation ($\pi^e_{i,t}$), the output gap and a temporary supply shock ($\epsilon_{2,i,t}$):

$$\pi_{i,t} = \pi^e_{i,t} + \beta y_{i,t} + \epsilon_{2,i,t}, \qquad \beta > 0 \tag{2}$$

In a standard closed-economy three-equation model, the model would be closed with a *MP* rule, which spells out how the central bank sets the real interest rate in response to developments in inflation and output. In our setting, we replace the *MP* rule by the following expression for the expected real interest rate:

$$r_{i,t} = i_{s,t} - \pi^e_{i,t} + \rho_{i,t}. \tag{3}$$

In equation (3), $i_{s,t}$ is the nominal short-term interest rate. In contrast to the standard model, equation (3) now includes the term $\rho_{i,t}$, which is a country-specific risk premium. The inclusion of $\rho_{i,t}$ serves to illustrate the effect of fragmentation risk on the economies of EA countries.[2] Although equation (3) does not constitute a traditional *MP* rule, all three ingredients are related to the

[1] This is similar to the critique of Kleidon (1986) in the debate on excess stock market volatility, in which he claims that fundamental prices constructed with ex-post data differ from market forecasts made under uncertainty.

[2] In a similar way, Stevenson and Wolfers (2023) add financial shocks to the *IS-MP-PC*-model.

monetary policy of the ECB. The policy rate set by the ECB is $i_{s,t}$. A credible monetary policy would anchor inflation expectations to the ECB's objective, setting $\pi^e_{i,t}$ equal to the inflation target π_T. Finally, $\rho_{i,t}$ measures the fragmentation in interest rates for which the ECB has introduced the new TPI instrument.[3]

In this model, interest-rate policy is of little help in macroeconomic stabilization. Given the small-country assumption, $i_{s,t}$ will not react to output and inflation in country i. For example, a negative asymmetric demand shock in country i will shift the *IS* curve to the left, thereby reducing the output gap. Via the Phillips curve in equation (2), this will also lower inflation in country i. The policy rate in equation (3) will, however, not be reduced to stabilize the economy and move output and inflation back to their original positions. If anything, equation (3) points to two forces which may destabilize the economy of country i.

First, in the absence of a stabilizing adjustment mechanism, output and inflation in country i may remain low. This risks the de-anchoring of inflation expectations from the ECB's target ($\pi^e_{i,t} < \pi_T$). If that were to happen, the Phillips curve in equation (2) would shift downwards and the expected real interest rate in equation (3) would shift upwards. Together, these two effects would amplify the disinflationary contraction in country i. Analytical support for this procyclical effect of real interest rate divergence in a monetary union is provided by Bofinger and Mayer (2007).[4] Without countervailing forces, this process of macroeconomic destabilization may persist. Note that any regional de-anchoring of inflation expectations gives rise to real interest rate differentials across EA countries, even with a uniform nominal policy rate and zero risk premia. This implies that in the absence of bond market fragmentation, real lending conditions for firms and households can still diverge across the union.

A second force which may destabilize country i is negative bond market sentiment. As discussed above, the possibility that unwarranted disorderly market dynamics fragment the monetary union has been the reason why the ECB has introduced the TPI. In our model, negative market sentiment manifests itself in investors demanding a positive risk premium $\rho_{i,t}$. Via the effect on $r_{i,t}$, output in country i decreases along the IS curve, resulting in disinflationary pressure via the Phillips curve. This contractionary effect of negative market sentiment will be amplified when inflation expectations become de-anchored, leading to a further increase in $r_{i,t}$ and a downwards shift in the Phillips curve. Moreover, a positive feedback loop may develop, whereby the drop in output heightens bond market concerns about the creditworthiness of country i and further increases $\rho_{i,t}$. The latter effect will depend on the state of the public finances and the likelihood that country i falls into a debt trap, whereby economic contraction and increased interest expenses project an unsustainable trajectory for the debt-to-GDP ratio. By activating the TPI, the ECB could thwart this self-fulfilling destabilizing effect of financial fragmentation by intervening in the bond markets to reduce the risk premium. In effect, with the TPI the ECB has introduced an instrument which works on the level of individual countries, in addition to the union-wide short-term nominal policy rate. A further procyclical effect, not included in equation (3), works through the housing market. Divergences in economic growth and real interest rates may aggravate fragmentation in EA housing markets and lead to procyclical wealth effects.

We finally discuss potential countervailing forces, which, in the absence of a monetary policy response, may halt the process of macroeconomic destabilization. In principle, countercyclical national fiscal policy could be used for stabilization purposes, by countering economic contraction with fiscal stimulus, shifting the *IS* curve back to right. This option may, however, not be available in an environment of negative bond market sentiment, when investors question the creditworthiness of the sovereign. Instead, cross-country fiscal transfers could be used, shifting resources from booming to depressed regions within the union. However, the fiscal capacity of the EA is still quite limited. Even when sizable resources are made available, as in the Next Generation EU programme, the question remains whether this funding is useful for short-run stabilization purposes, given the delays in decision-making and execution.

This leaves the real exchange rate as the main adjustment channel (Bofinger and Mayer 2007). The contraction in country i will improve the competitive position through a process of internal devaluation. This will be reflected in an increase in net exports, shifting the *IS* curve back to the right. In contrast to the adjustment via the nominal exchange rate, internal devaluation can be a slow process, as the sovereign debt crisis has shown (see also Arnold and Kool 2004).

Following Goodhart and Hofmann (2005a) and Goodhart and Hofmann (2005b), equation (4) shows an extended *IS* curve, adding wealth and real exchange rate channels to the real interest rate channel:

$$y_{i,t} = \alpha_0 - \alpha_1 r_{i,t} - \alpha_2 rex_{i,t} - \alpha_3 hp_{i,t} + \epsilon_{1,i,t}, \qquad \alpha_1 > 0, \alpha_2 > 0, \alpha_3 > 0 \tag{4}$$

In equation (4), $rex_{i,t}$ and $hp_{i,t}$ denote respectively the real exchange and the housing price level in country i. This equation will be the basis for our empirical specification.

3. Data and method

3.1. Data

Our dataset comprises quarterly time series from 1999Q1 to 2023Q3 for a panel of twelve EA countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain. For some variables, the data start later. See Table A.4 in appendix A for further details on sources and sample periods.

[3] As the *IS-MP-PC* model does not incorporate the maturity spectrum of interest rates, equation (3) combines factors which influence the short end ($i_{s,t}$) and the long end ($\rho_{i,t}$) of the yield curve. We will disentangle these effects in the empirical specification.

[4] This mechanism is also akin to the Walters critique of pegged exchange rates in the presence of diverging inflation rates, see Walters (1990).

We measure the output gap of country i, denoted $y_{i,t}$, as the percentage gap between real GDP and potential real GDP. GDP data are taken from the OECD database. We calculate potential real GDP using a Hodrick-Prescott filter with a smoothing parameter of 1600, cf. Goodhart and Hofmann (2005a). To account for the sudden and sizeable drop in output due to the outbreak of the pandemic we include the dummy variable $covid$, which takes on the value of one in the first two quarters of 2020 and zero otherwise.

We decompose the expected real interest rate ($r_{i,t}$) into the expected risk-free real short-term rate, denoted $r_{s,i,t}$, and a risk premium ($\rho_{i,t}$). We measure $r_{s,i,t}$ as follows. First, we construct a series for the nominal short-term rate ($i_{s,t}$) in the EA. During most of the sample period, short-term rates have been at the zero lower bound. We therefore use the EA shadow rate constructed by Wu and Xia (2017). As the shadow rate runs from 2004Q3 to 2022Q2, we use money market rates for the periods before and after this time window. We construct national measures of one-year ahead expected inflation using data from the European Commission's Consumer Survey. The use of Consumer Survey data has the advantage that it provides a direct measure of consumers' inflation expectations based on a large-scale survey, in contrast to measures based on yield curves or small-scale surveys among professional economists. A further advantage is that the data are available for all EA countries. The main limitations of these data are that the forecast horizon is one year and that the data are qualitative. The former limitation is not a problem in our setting, as we are measuring the expected real short-term rate. The latter limitation is resolved by extracting quantitative inflation expectations from the qualitative data. Appendix B briefly summarizes the methodology that is used. The expected risk-free real short-term rate is next determined by subtracting the expected inflation from the nominal short-term rate.

For the extended *IS* curve, we include real housing prices, denoted $hp_{i,t}$. Data on housing prices have been taken from the BIS database and are deflated using HICP inflation excluding energy and food prices. We also include the real effective exchange rate, denoted $rex_{i,t}$. The data have been taken from the BIS database and measure the competitiveness of country i vis-à-vis a basket of 64 economies.

Risk premia due to fragmentation risk are measured using interest rate spreads. The variable $i10y_{spr,i,t}$ measures nominal sovereign spreads, defined as the difference between the yield on 10-year government bonds of country i and the average yield of the EA countries in our sample. To measure fragmentation in private lending conditions, we use the composite cost of borrowing indicators from the ECB. These indicators are based on the ECB's MFI interest rate statistics and are constructed to enhance the comparability of borrowing costs for non-financial corporations and households across EA countries. The ECB distinguishes the following four categories (the variable names of the corresponding spreads vis-à-vis the EA average are in parentheses):

1. Cost of borrowing indicator for households for house purchase ($cobhh_{spr,i,t}$);
2. Cost of borrowing indicator for non-financial corporations ($cobnfc_{spr,i,t}$);
3. Cost of borrowing indicator for short-term loans to households and non-financial corporations ($cobst_{spr,i,t}$);
4. Cost of borrowing indicator for long-term loans to households and non-financial corporations ($coblt_{spr,i,t}$).

3.2. Method

In most simple business cycle models, a key role is played by a dynamic *IS* curve relating the output gap to a measure of the real interest rate. A sizable empirical literature has developed in which various empirical specifications have been estimated, see e.g. Rudebusch and Svensson (1999), Peersman and Smets (1999), Goodhart and Hofmann (2005a), Goodhart and Hofmann (2005b) and Hafer and Jones (2008). Stracca (2017), Hawkins and Nguyen (2018) and van den End et al. (2020) provide recent additions to this literature. A major issue in this literature is the so-called *IS* puzzle, which refers to the failure of some studies to find a significant effect of the real interest rate on the output gap (Nelson 2002). Based on the model in section 2.2, we estimate an extended *IS* curve using a panel regression model with the following specification:

$$y_{i,t} = \alpha_i + \beta_1 y_{i,t-1} + \beta_2 y_{i,t-2} + \beta_3 covid + \beta_4 r_{s,i,t-1} + \beta_5 i_{spr,i,t-1} + \beta_6 rex_{i,t-1} + \beta_7 hp_{i,t-1} \tag{5}$$

The specification includes two lags for the output gap and the *covid* dummy variable to account for the unexpected output drop in the first two quarters of 2020. In contrast to equation (4), we disentangle the expected real interest rate into the expected risk-free real short-term rate ($r_{s,i,t}$) and the interest rate spread $i_{spr,i,t}$. For $i_{spr,i,t}$, we will use both nominal and real spreads for 10-year government bonds and for the cost of borrowing indicators listed in section 3.1. Due to multicollinearity, these spreads will be entered into the regression model separately. For both $r_{s,i,t}$ and $i_{spr,i,t}$ we use a one-quarter lag, which is in line with the literature. The variables $hp_{i,t}$ and $rex_{i,t}$ are included in lagged year-on-year log differences, due to the non-stationarity of the levels of these variables. Equation (5) also includes fixed country effects.

The panel model is estimated using GLS with cross-section weights, to account for cross-section heteroskedasticity. Standard errors are calculated using a two-way cluster robust coefficient covariance, which is robust to contemporaneous correlation between cross-section units and period correlation within cross-sections. Using panel unit root tests, we found that almost all variables are stationary, including the residuals from the panel model. As a further check, we have also estimated all specifications using the DOLS panel approach (Stock and Watson 1993). The DOLS approach accommodates heterogeneity in the short-run dynamics by including leads and lags of first differenced explanatory variables. As this did not change the results, these estimates go unreported.

As a next step, we combine the effects of monetary policy and fragmentation into a Monetary Conditions Index (MCI). An MCI is a weighted average of variables representing the impact of monetary policy, expressed as deviations from their values in a base year. The weights are based on the relative effect of these variables on output. Ericsson et al. (1997) and Costa (2000) provide an introduction to the subject. MCI's have been frequently used by monetary policymakers. Arguments in favour of using MCI's are simplicity and convenience. By including multiple monetary indicators a MCI offers an improvement over using the interest rate as the sole indicator of the monetary policy stance.

Originally, the MCI has measured the combined effects of changes in interest rates and exchange rates on output. However, since the introduction of the euro, the EA has become a more closed economy and the effect of short-term exchange rate variability on GDP has weakened substantially (Peeters 1999). In other words, as the openness of the region has declined, the information value of the exchange rate for monetary policy purposes has been reduced. In parallel, part of the economic uncertainty that in the pre-euro era led to exchange rate volatility, now manifests itself in the bond markets, resulting in fragmentation risk. In the context of a more closed but also fragmented monetary union, it therefore makes sense to replace the exchange rate with a measure of fragmentation risk in the MCI. Based on our extended *IS* curve estimates, an MCI can be constructed for each EA country i as follows:

$$mci_{i,t} = \beta_4(r_{s,i,t} - r_{s,i,0}) + \beta_5(i_{spr,i,t} - i_{spr,i,0}) \tag{6}$$

In equation (6), the deviations of $r_{s,i,t}$ and $i_{spr,i,t}$ from their base values $r_{s,i,0}$ and $i_{spr,i,0}$ are weighted by the corresponding regression coefficients in equation (5), which measure their effects on the output gap. For presentation purposes, we rescale $mci_{i,t}$ to an index which equals 100 in the base year and which increases (decreases) when monetary conditions tighten (loosen):

$$MCI_{i,t} = MCI_0 exp[|\beta_4|(r_{s,i,t} - r_{s,i,0}) + |\beta_6|(i_{spr,i,t} - i_{spr,i,0})] \tag{7}$$

In equation (7), MCI_0 equals 100 and the β's are taken in absolute values. Below, we plot the MCI's of EA countries based on estimates of the extended *IS* curve. With our MCI estimates, we calculate two indicators of fragmentation risk. The first one is σ-convergence, which focuses on the cross-sectional dispersion in MCI's across EA countries. Absent fragmentation, σ-convergence would be zero. A drawback of this measure is that it doesn't show whether cross-country differences in monetary conditions are temporary or persistent. We therefore add a measure of β-convergence, which is estimated by regressing the change in the MCI on the lagged level of the MCI:

$$\Delta MCI_{i,t} = \alpha_i + \gamma_t + \beta MCI_{i,t-1} \tag{8}$$

Equation (8) is estimated using OLS and includes fixed country effects α_i and fixed period effects γ_t. Standard errors are calculated using a two-way cluster robust coefficient covariance. The more negative β, the higher the convergence speed. We also calculate the speed of convergence λ and the half-life τ (cf. Arbia and Baltagi 2008).

Notwithstanding the fact that our approach to estimating the *IS* curve is quite common in the empirical literature, it is not without flaws. A common criticism is that, due to the forward-looking nature of monetary policy, the reduced-form specification in equation (5) suffers from simultaneity bias. Because of this, many studies on the effects of monetary policy employ a VAR approach to identify the exogenous or unsystematic component of monetary policy. However, as argued in Goodhart and Hofmann (2005a), a drawback in using a VAR is the focus on monetary policy shocks. These shocks may constitute just a part of all movements in real interest rates and do not shed light on the systematic effects of monetary policy. This point is relevant in the EA context. Persistent inflation differentials across EA countries in the presence of a uniform nominal policy rate may generate systematic differences in real rates, cf equation (3). As a robustness check, however, we examine the output effect of our main fragmentation variable using a five-variable panel VAR including y_i, π_i, hp_i, $r_{s,i}$, rex_i and $i10y_{spr,i}$. We calculate the impulse response of the output gap to an innovation in $i10y_{spr,i}$ and compare this to the corresponding coefficient estimates of the extended *IS* curve.

4. Results

4.1. IS curve estimation

Table 1 reports the results of estimating various specifications of the extended IS curve in equation (5). In all specifications, the lagged output gap and the covid dummy variable are significant at at least a 5% level. Column (a) shows the benchmark specification including the lagged expected risk-free real short-term rate. The coefficient of $r_{s,i,t-1}$ is significant at a 5% level and implies that a one percentage point increase in $r_{s,i,t-1}$ reduces the output gap by 0.06%. This estimate is in the range of estimates in Goodhart and Hofmann (2005a). As a robustness check, we have also estimated specification (a) with a backward-looking real interest rate, by replacing expected inflation with lagged inflation. This didn't materially change the results. We add the real exchange rate in specification (b). This variable is marginally significant, which is in line with the observation that the EA has come to resemble a closed economy. Specification (c) adds the spreads in 10-year sovereign bond yields. The coefficient of $i10y_{spr,i,t}$ is highly significant at a 1% level. The value of the coefficient (-0.098) implies that a one percentage point increase in the spread has a stronger effect on the output gap than a similar increase in the expected real short-term rate. Specification (d) finally includes the effect of housing prices on the output gap. Due to the data availability of housing prices, the number of observations is somewhat lower. As in Goodhart and Hofmann (2005a) and Goodhart and Hofmann (2005b), the coefficient of $hp_{i,t-1}$ is positive and highly significant. At the same time, the inclusion of $hp_{i,t-1}$ reduces the coefficient of $i10y_{spr,i,t}$ from -0.098 to -0.076. This raises the question which of the two coefficients should be used as an input for the MCI calculation in equation (7). It can be argued that one of the channels through which financial fragmentation affects the economy is the housing market. Fragmentation in long-term bond yields may translate into fragmentation in mortgage rates and affect the output gap through housing prices. Following this line of reasoning, the coefficient of $i10_{spr,i,t}$ in specification (d) would underestimate the effect of financial fragmentation, as the transmission via the housing market is captured by $hp_{i,t-1}$. For this reason, we will use the coefficients of specification (c) in constructing the MCI's.

Table 1
Extended IS curve.

	(a)	(b)	(c)	(d)
$y_{i,t\text{-}1}$	0.576***	0.573***	0.567***	0.536***
	(0.173)	(0.173)	(0.169)	(0.161)
$y_{i,t\text{-}2}$	0.153	0.152 0.149	0.159	
	(0.111)	(0.110)	(0.107)	(0.101)
covid	-7.609**	-7.651**	-7.679**	-7.591**
	(3.201)	(3.19)	(3.218)	(3.208)
$r_{s,i,t\text{-}1}$	-0.059**	-0.059**	-0.058**	-0.0521**
	(0.022)	(0.022)	(0.022)	(0.023)
$rex_{i,t\text{-}1}$	-0.036*	-0.039*	-0.042**	
		(0.019)	(0.018)	(0.017)
$i10y_{spr,i,t\text{-}1}$			-0.098***	-0.076**
			(0.023)	(0.028)
$hp_{i,t\text{-}1}$				0.050***
				(0.011)
Adjusted R^2	0.64	0.64	0.65	0.66
DW	2.14	2.14	2.16	2.24
N	1116	1116	1116	1016

Note: ***, ** and * denote significance at a level of 1%, 5% and 10%. Standard errors are in parentheses. Sample period: 2000Q2-2023Q2.

Table 2
Coefficients of fragmentation measures.

	$i10y_{spr,i,t\text{-}1}$	$cobhh_{spr,i,t\text{-}1}$	$cobnfc_{spr,i,t\text{-}1}$	$cobst_{spr,i,t\text{-}1}$	$coblt_{spr,i,t\text{-}1}$
nominal	-0.098***	-0.130	-0.150	-0.155	-0.084***
	(0.023)	(0.121)	(0.167)	(0.167)	(0.022)
real	-0.048*	0.034*	0.026	0.026	0.022
	(0.025)	(0.017)	(0.032)	(0.033)	(0.046)

Note: *** and * denote significance at a level of 1% and 10%. Standard errors are in parentheses.
Sample period: 2000Q2-2023Q2; for cost of borrowing spreads: 2003Q02-2023Q2.

We next estimate different versions of specification (c) by separately entering both nominal and real spreads for 10-year government bonds and for private sector cost of borrowing indicators into the regression model. Table 2 summarizes the findings, by reporting the coefficients of the different spreads and their statistical significance. For all nominal spreads, the coefficients have the expected negative sign. This is not the case for the real spreads, for which only the coefficient of $i10y_{spr,i,t-1}$ is negative. The insignificance of the real spreads can be explained by the fact that $r_{s,i,t}$ may pick up the effect of cross-country variation in expected inflation. Among the spreads, $i10y_{spr,i,t-1}$ stands out. Its coefficient is significant at a 1% level for the nominal version, and at a 10% level for the real version. Among the cost of borrowing spreads, the coefficient of $coblt_{spr,i,t-1}$ is the only one that is negative and significant. Based on the estimates in Table 2, we will proceed by constructing MCI's for the two significant nominal fragmentation measures: $i10y_{spr,i,t}$ and $coblt_{spr,i,t}$.

4.2. Monetary conditions indices

Fig. 1 plots the MCI's based on equation (7) using nominal $i10y_{spr,i,t}$ as fragmentation measure (denoted MCI_{i10y}). In the interest of graph readability, the plot leaves out Greece, for which the MCI_{i10y} reached a value of 500 in early 2012.[5] Fig. 1 shows the effect of the sovereign debt crisis on monetary conditions across the EA. The MCI_{i10y}'s of the smaller peripheral countries Ireland and Portugal strongly increase between 2010 and 2012, under the influence of rising sovereign spreads. During this period, a gap also emerges between the MCI_{i10y}'s of Italy and Spain and the rest of the EA countries. From 2012 to 2021, the loosening of monetary policy using unconventional instruments reduces the levels of and dispersion between the MCI_{i10y}'s. From 2022, the tightening of monetary policy is reflected in the increase in the MCI_{i10y}'s.

Fig. 2 plots the MCI's using nominal $coblt_{spr,i,t}$ as our fragmentation measure (denoted MCI_{coblt}). Again we exclude Greece, this time to enable a proper comparison with Fig. 1. While Fig. 2 also points to cross-country dispersion in MCI_{coblt}, compared to Fig. 1 the fragmentation during the sovereign debt crisis is much less pronounced. In contrast, the recent surge in the MCI_{coblt}'s is similar to that in Fig. 1.

[5] The extreme values of Greek bond yields during the period from 2010 to 2015 have a large influence on graphs and statistical measures. While one could argue that the Greek case is a fitting illustration of fragmentation risk, we want to avoid that the empirical results depend on extreme observations. The inclusion of Greece would, however, have strengthened our empirical findings on fragmentation.

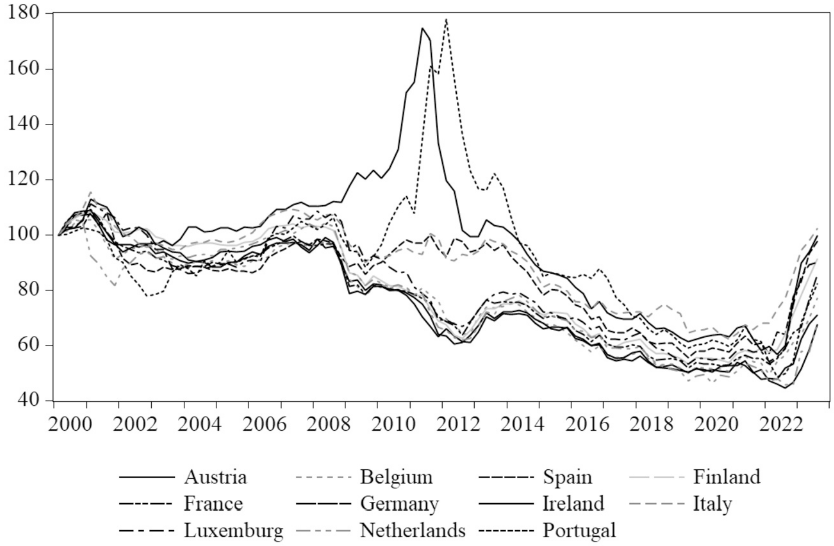

Fig. 1. MCIs including sovereign spreads.

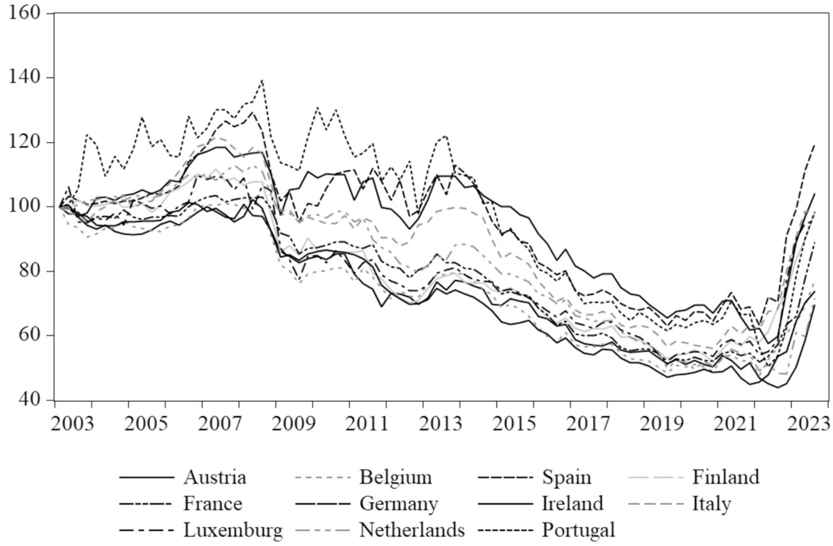

Fig. 2. MCIs including spreads in long-term cost of borrowing.

We finally compare the evolution of MCI_{i10y} and MCI_{coblt} for two groups of countries: a group consisting of distressed EA countries (Italy, Ireland, Portugal & Spain, but excluding Greece, denoted IIPS) and a group with other EA countries (Austria, Belgium, Germany, Finland, France, Luxemburg & The Netherlands, denoted non-IIPS). For both groups, unweighted averages of the MCI's are constructed. Fig. 3 shows that since 2009 the MCI's for the IIPS group have consistently been higher than for the non-IIPS group. While the MCI_{i10y}'s for the two groups were very close before the Global Financial Crisis, possibly reflecting unwarranted bond market complacency about sovereign risk, the two series start to diverge strongly during the crisis, before converging from 2013 to 2015. Notwithstanding the ECB's unconventional monetary policy measures, a gap between the MCI_{i10y}'s of the two groups has remained. Regarding MCI_{coblt}, the graph shows that monetary conditions were more tight in the IIPS group even before the crisis. While the effect of the crisis on MCI_{coblt} is less pronounced, Fig. 3 shows a persistent gap between the MCI_{coblt}'s of the two groups. Fig. 3 suggests that, while the strong increase in the difference in MCI_{i10y} between IIPS and non-IIPS countries is crisis-related, more moderate levels of fragmentation have been a persistent feature of the EA.

4.3. Convergence measures

Fig. 4 shows the (lack of) σ-convergence in MCI_{i10y} and MCI_{coblt}. The σ's are calculated as the cross-sectional dispersion in MCI's across EA countries, weighted by real GDP and excluding Greece. There is no indication that the σ's have gone down over

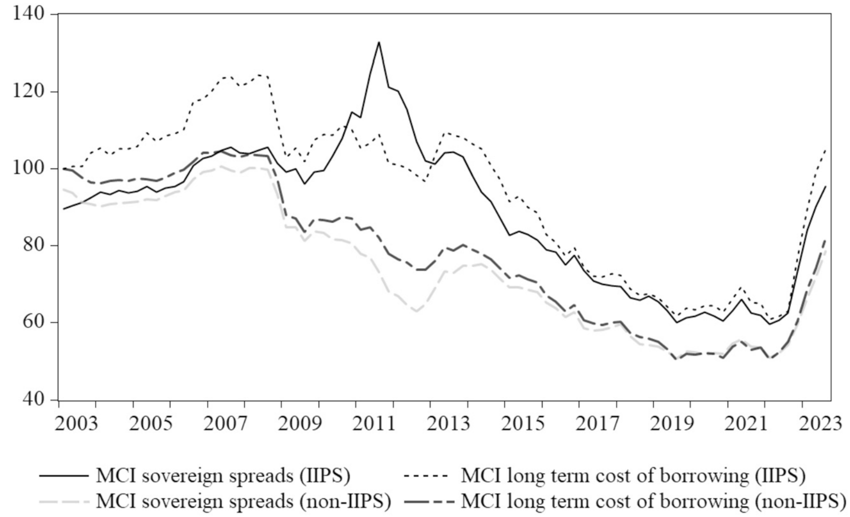

Fig. 3. IIPS versus non-IIPS countries.

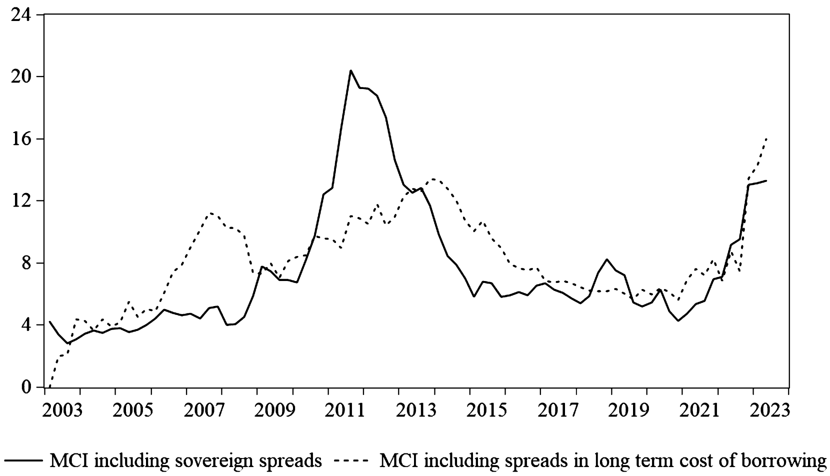

Fig. 4. Dispersion in MCIs.

time. In fact, fitting a linear trend to the dispersion in MCI_{i10y} and MCI_{coblt} in both cases yields a positive and significant trend coefficient.[6] While the dispersion in MCI_{coblt} is elevated during the sovereign debt crisis, the increase did not match the spike in the dispersion in MCI_{i10y}. Also, the recent dispersion in MCI_{coblt} is at a similar level as during the sovereign debt crisis, suggesting that the EA economy should be able to cope with such a level of fragmentation. The relatively minor fluctuations in the dispersion of MCI_{coblt} suggest that impediments to the uniform transmission of monetary policy to private lending rates have not been the main source of fragmentation risk in the EA. Rather, Fig. 4 suggests that during the euro crisis, confidence crises in EA bond markets and subsequent austerity measures in distressed countries may have been the main driver of fragmentation risk. In our extending *IS* curve, the effect of panic-driven fiscal adjustment will be picked up by $i10y_{spr,i,t}$. Because of this, one could argue that MCI_{i10y} partly captures fiscal instead of monetary conditions. In this line of reasoning, the MCI_{coblt}'s better capture the non-uniform transmission of monetary policy to private lending conditions, as this metric will be less contaminated with the effect of a crisis-induced fiscal response on output.

To facilitate the interpretation of Fig. 4, Fig. 5 shows the cross-sectional dispersion in $i10y_{spr,i,t}$, $coblt_{spr,i,t}$ and $r_{s,i,t}$. The evidence of bond market fragmentation is strong, as illustrated by the peak in the cross-sectional dispersion in bond yields during the sovereign debt crisis. The peak in bond yield dispersion is followed by a clear hump in the dispersion of $coblt_{spr,i,t}$. Finally, Fig. 5 shows the strong surge in the dispersion of $r_{s,i,t}$ in the post-covid period, when the dynamics of high inflation dominates the dispersion in real interest rates. Comparing Figs. 5 and 4 leads to the conclusion that the recent increase in the dispersion in MCI_{i10y} and MCI_{coblt}

[6] The trend coefficient is 0.041 (p-value 0.034) for the dispersion in MCI_{i10y} and 0.038 (p-value 0.005) for the dispersion in MCI_{coblt}.

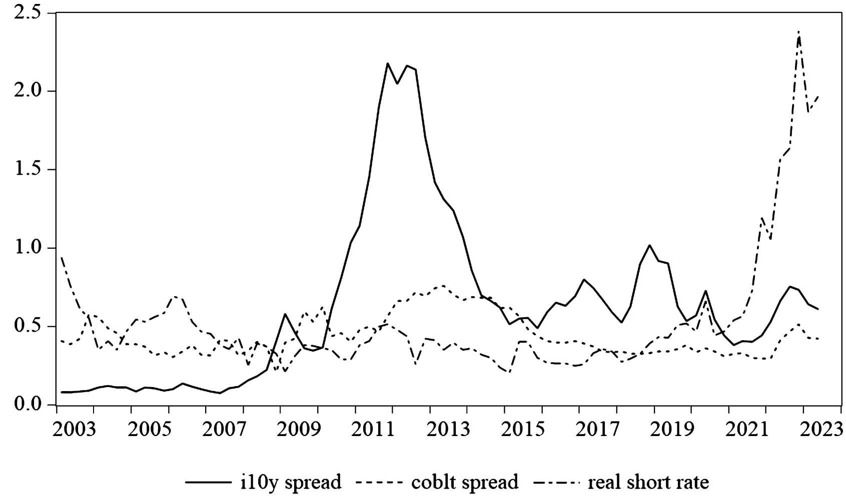

Fig. 5. Dispersion in interest rates.

Table 3
β convergence.

	Including Greece			Excluding Greece		
	β	λ	τ	β	λ	τ
MCI_{i10y}	-0.047	0.048	14.43	-0.059***	0.061	11.44
	(0.028)			(0.013)		
MCI_{coblt}	-0.098***	0.103	6.72	-0.117***	0.125	5.55
	(0.019)			(0.023)		

Note: *** denotes significance at a level of 1%.
Standard errors are in parentheses. Sample period: 2003:1-2023:2.

is triggered by an increase in the dispersion of the real short-term rates, whereas the increase in the dispersion in the MCI's during the sovereign debt crisis is related to an increase in the dispersion of spreads.

As Fig. 4 is silent on the persistence of cross-country differences in MCI's, Table 3 reports our estimates of β-convergence in MCI_{i10y} and MCI_{coblt}. We report estimates both including and excluding Greece, to show the effect of this extreme observation. Two observations stand out. First, the β-estimates are lower when we include Greece. Including Greece, the β coefficient for MCI_{i10y} even becomes insignificant. Second, β-convergence is much stronger for the MCI_{coblt}'s, suggesting that fragmentation in private lending conditions may be less of a problem than fragmentation in sovereign bond yields.

4.4. *Robustness check*

We next examine whether the use of a VAR approach leads to a different effect of bond market fragmentation on output. The panel VAR includes the variables y_i, π_i, hp_i, $r_{s,i}$, rex_i and $i10y_{spr,i}$. The VAR includes fixed effects and the *covid* dummy variable for output. The identification of shocks is done with a standard Cholesky factorisation, using the following ordering: y_i, π_i, hp_i, $r_{s,i}$, rex_i and $i10y_{spr,i}$. For the most part, this ordering is standard in the VAR literature, see e.g. Goodhart and Hofmann (2001). The ordering assumes that the output gap may have a contemporaneous effect on inflation. Both y_i and π_i do not contemporaneously react to the other variables. As housing prices are presumed to be more sticky than other financial variables, they are ranked third. In contrast, bond prices are flexible. We assume that price formation in EA bond markets reacts contemporaneously to all other variables. Therefore, $i10y_{spr,i}$ is ranked last. Regarding the ordering of $r_{s,i}$ and rex_i, we have put $r_{s,i}$ before rex_i, as the EA does not follow an exchange rate policy. Changing the order between these two variables did not, however, affect the results.

Fig. 6 shows the impulse responses of the output gap to a one standard deviation innovation in the sovereign spread over a period of twelve quarters after the shock. The dotted lines indicate the 95% confidence interval. The impulse responses are negative and become significant at a 5% level after three quarters. To compare these estimates to the results from the extended *IS* curve estimation, we follow Goodhart and Hofmann (2001) in averaging the impact of the innovation in the spread over twelve quarters. The average impact is -0.154 for a one standard deviation shock to $i10y_{spr,i}$. To compare this number to the coefficients of $i10y_{spr,i}$ in columns (c) and (d) of Table 1, we multiply these by the standard deviation in $i10y_{spr,i}$, which is 1.9. The resulting output effects are respectively -0.186 and -0.144, which are in the same order of magnitude as the average impulse response.

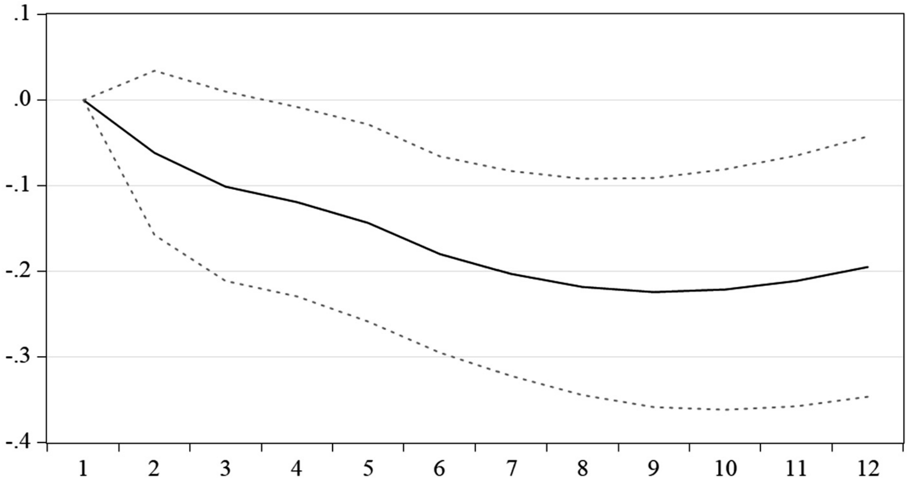

Fig. 6. Response of output gap to innovation in i10y spread.

5. Conclusions

In March 2020, during a press conference at the start of the pandemic, ECB president Lagarde commented that the ECB is "not here to close spreads" (FT 2020). Two years later, following the normalization of monetary policy and the resulting increase in sovereign spreads between EA countries, the ECB introduced the TPI, the new policy instrument for spread control. This policy shift illustrates that an active role of the ECB in reducing fragmentation risk is not self-evident and requires a sound rationale. A compelling case for policy interventions aimed at mitigating financial fragmentation rests on non-trivial output effects. As the primary concern with fragmentation lies in its potential to trigger disparities in output and inflation among EA countries, any assessment of fragmentation risk should move beyond calculating the dispersion in interest rates and also factor in their impact on output.

Academic research in this area is still scarce. The current paper contributes to the literature in the following ways. First, we estimate the output effects of financial fragmentation using an extended IS curve, incorporating various fragmentation measures. As a second contribution, this paper proposes consolidating the joint effects of monetary policy and financial fragmentation into a MCI for EA countries. This index may serve as a more comprehensive measure of monetary conditions in a fragmented monetary union, capturing not only the magnitude of policy rates and spreads but also their impact on output.

Our estimates of an extended IS curve for a panel of twelve EA countries show a significant negative effect of the expected risk-free short-term real interest on the output gap. In addition, we find that sovereign spreads and spreads in the long-term borrowing costs of the private sector have a sizable and significant effect on output. This finding supports the ECB's stance that financial fragmentation can impede a uniform transmission of monetary policy to the real EA economy. In contrast, the output effect of the real exchange rate is weak, confirming the reduced importance of exchange rate movements in the EA and justifying the replacement of the traditional role of the exchange rate in a MCI with fragmentation measures.

Based on the IS curve estimates, we construct two versions of the MCI for all EA countries, using respectively sovereign spreads and spreads in the long-term cost of borrowing. These fragmentation-enhanced measures of the monetary policy stance show considerable cross-country variation. We look at σ-convergence and β-convergence of our MCI estimates. The σ's of both MCI versions show no sign of convergence. If anything, they trend upward. During the sovereign debt crisis, the cross-sectional dispersion of MCI's based on sovereign spreads peaked and exceeded those based on the private cost of borrowing. More recently, the σ's of both MCI's have surged as the post-covid inflation wave has increased the dispersion in real interest rates. With regard to β-convergence, we find that the MCI's based on sovereign spreads converge at a much slower speed than the MCI's based on the private cost of borrowing.

This paper suggests that the cause of fragmentation in monetary conditions, and thus the appropriate policy response, may change over time. Insofar as the current fragmentation is related to the recent outburst of inflation, a clear role for the ECB is to restore an environment of low and stable inflation. Insofar as heightened fragmentation risk has a fiscal origin, as during the sovereign debt crisis, one could argue that a fiscal policy response, either national or union-wide, is more appropriate than a monetary policy response. A further finding in this paper is that during the crisis the fragmentation in MCI's based on the private cost of borrowing did not rise to the level of the fragmentation in MCI's based on sovereign spreads. Combined with the stronger β-convergence in MCI's based on the private cost of borrowing, this suggests that a non-uniform monetary policy transmission to private lending conditions is not the most important driver of fragmentation risk. Nevertheless, our results show a persistent gap between monetary conditions in IIPS and non-IIPS countries, also for the MCI's based on the private cost of borrowing. We conclude that, while there are strong reasons for the ECB to be concerned about the fragmentation in monetary conditions across the EA, this does not necessarily imply that the TPI is the best instrument to deal with it.

Table A.4
Data.

variable	source	frequency	period
real GDP	OECD	quarterly	1999Q1-2023Q2
housing prices	BIS	quarterly	1999Q1-2023Q2 2000Q1-2023Q2 (AT) 2006Q1-2023Q2 (GR) 2007Q1-2023Q2 (LU) 20081-2023Q2 (PT)
real effective exchange rate	BIS	monthly	1999M1-2023M6
shadow rate	https://sites.google.com/view/jingcynthiawu/shadow-rates	monthly	2004M9-2022M8
call interbank rate	FRED	daily	1/1/1999-30/6/2023
10-year government bond yields	Eurostat	monthly	1999M1-2023M6
cost of borrowing indicators	ECB	monthly	2003M1-2023M6
qualitative inflation expectations	European Commission	monthly	1999M1-2023M6
HCIP inflation ex. energy and food	Eurostat	monthly	1999M1-2023M6

CRediT authorship contribution statement

Ivo J.M. Arnold: Writing – original draft, Resources, Methodology, Investigation, Formal analysis, Data curation, Conceptualization.

Appendix A. Data appendix

Table A.4 list the sources, sample periods and frequencies of the data used in this paper. All higher-frequency data have been converted to quarterly data by averaging.

Appendix B. Construction of inflation expectations

In the Consumer Survey of the European Commission, respondents are asked about their expectations regarding the development of consumer prices. We use the responses to Questions [5] and [6]. Question [5] asks consumers to assess price developments over the past year: "How do you think that consumer prices have developed over the last 12 months? They have. . .

1. Risen a lot
2. Risen moderately
3. Risen slightly
4. Stayed about the same
5. Fallen
6. Don't know".

Question [6] asks consumers about future price developments: "By comparison with the past 12 months, how do you expect consumer prices will develop in the next 12 months? They will . . .

1. Increase more rapidly
2. Increase at the same rate
3. Increase at a slower rate
4. Stay about the same
5. Fall
6. Don't know".

The literature on the extraction of quantitative inflation expectations from qualitative survey responses uses the so-called probability approach, according to which the shares of responses in each response category can be interpreted as estimates of areas under the density function of aggregate inflation expectations (i.e., as a probability). Forsells and Kenny (2003) provide a methodological exposition. The probability approach requires the specification of a distribution function. In line with much of the literature, we use a logistic distribution.

The extraction procedure requires a measure for perceived inflation. Perceived inflation (π_t^p) is derived from the survey response to Question [5] as follows, cf. Dias et al. (2010):

$$\pi_t^p = -\pi_{ex,t}\left(\frac{Z_t^3 + Z_t^4}{Z_t^3 + Z_t^4 - Z_t^3 - Z_t^4}\right) \tag{B.1}$$

where the Z_t^i's in equation (B.1) reflect the statistical distribution of Question [5] in the Consumer Survey. They are determined as follows:

$$\begin{aligned} Z_t^1 &= N^{-1}[1 - S_t^1], \\ Z_t^2 &= N^{-1}[1 - S_t^1 - S_t^2], \\ Z_t^3 &= N^{-1}[1 - S_t^1 - S_t^2 - S_t^3], \\ Z_t^4 &= N^{-1}[S_t^5], \end{aligned} \tag{B.2}$$

where S_t^i is the sample proportion for response category i, and N^{-1} refers to the inverse of the cumulative logistic distribution function. As smoothed inflation measure to scale inflation perceptions we use the HCIP inflation excluding energy and food prices, denoted $\pi_{ex,t}$. We next use our measure for perceived inflation and the responses to Question [6] from the Consumer Survey to derive the mean expected inflation 12 months ahead (π_{t+12}^e):

$$\pi_{t+12}^e = -\pi_t^p \left(\frac{Z_t^3 + Z_t^4}{Z_t^1 + Z_t^2 - Z_t^3 - Z_t^4} \right), \tag{B.3}$$

where the Z_t^i's in equation (B.3) reflect the statistical distribution of Question [6] from the Consumer Survey and are determined according to equation (B.2). As Irish survey data were unavailable before 2016, we use regression analysis to estimate a relationship between inflation and expected inflation at the EA level and use the coefficient estimates to generate Irish inflation expectations.

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