Macroprudential Policies, Economic Growth and Banking Crises☆

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Abstract

Using a sample covering emerging market and advanced economies, we assess the impact of macroprudential policies on financial stability. Our empirical setup is designed to account for the potential direct and indirect effects that macroprudential policies can have on banking crises. We find that while macroprudential policies (MPPs) exert a direct stabilizing effect, they also have an indirect destabilizing effect, which works through the depressing of economic growth. It turns out that mitigating effects of MPPs on the likelihood of banking crises is more pronounced in emerging market economies relative to advanced economies.

Keywords: Macroprudential Policies, Banking Crises, Economic Growth

JEL Classification: C33, G01, G18

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1. Introduction

The Global Financial Crisis (GFC) provided a stark reminder of the dangers and costs of systemic financial crises, leading public authorities to increasingly rely on macroprudential policies to contain the risk of incidence of such crises. These policies, used to prevent the build-up of financial vulnerabilities and strengthen the resilience of the financial system to various types of shocks, have been in use for many years before the GFC, especially by emerging market economies (Cerutti et al., 2017a). Their use has, however, been stepped up after the GFC by both emerging market and advanced economies (Akinci and Olmstead-Rumsey, 2018).

Yet, our understanding of the extent to which macroprudential policies are effective in enhancing financial stability and lowering the incidence of financial crises, which is their ultimate goal, is far from being complete. In fact, to the best of our knowledge, apart from Choi et al. (2018), no previous studies have undertaken the task of systematically investigating the impact of macroprudential policies on the likelihood of financial crises. As we detail in the next section, the growing body of empirical research on macroprudential policies tends to emphasize the role that can be played by these policies in smoothing financial cycles by mitigating growth in credit and housing markets (e.g., Lim et al., 2011; Dell'Ariccia et al., 2012; IMF, 2012; Crowe et al., 2013; Claessens, 2014; Cerutti et al., 2017a). However, whether such mitigation of credit and housing market cycles eventually reduces the occurrence of financial crises remains an open question. In this paper, we depart from this approach and focus on the ultimate goal for which macroprudential policies are designed, which is making financial crises less likely. Specifically, we examine the impact of macroprudential policies on the likelihood of systemic banking crises.

Additionally, the effectiveness of macroprudential policies in fostering financial stability cannot be fully assessed by restricting the analysis to their credit and housing markets effects. Indeed, while there is increasing evidence pointing to the benefits of macroprudential policies with respect to smooth financial cycles, there are also growing concerns about their possible harmful effects on economic activity. In particular, by having a negative impact on credit supply and investment, macroprudential policies might depress economic growth. Using a panel of mostly OECD countries, Sanchez and Rohn (2016) report results suggesting that the use of macroprudential policies is costly as they are conducive to lower average economic growth. Kim and Mehrotra (2017) use a VAR framework and a sample of four countries from the Asia-Pacific region to analyze the responses of credit, output and inflation to changes in macroprudential and monetary policies. Their findings point to the existence of a negative impact of changes in macroprudential policies on economic output and inflation. Richter et al. (2019) quantify the effects of changes in maximum loan-to-value (LTV) ratios on output and "*find that over a four year horizon, a 10 percentage point decrease in the maximum LTV ratio leads to a 1.1 percent reduction in output*" (p.263). Any assessment of macroprudential policies should thus consider their potential damaging effect on the real economy. This is all the more important that weak economic growth can itself be a trigger of financial instability. Specifically, a slowdown in economic activity typically increases the amount of non performing loans and causes losses to the banking sector, which in the presence of vulnerabilities (excessive leverage, low liquidity buffers, etc.) can evolve into a full-blown crisis. Macroprudential policies whose aim is to ensure financial stability can thus end up being counterproductive by putting a drag on economic activity, which can in turn threaten financial stability. Despite the growing body of research pointing to the potential harmful effect of macroprudential policies on economic activity, no previous study has explored the implications of such an effect with respect to the ultimate goal of financial stability. In other words, we have no evidence on the net effect of macroprudential policies on the likelihood of financial crises.

This paper adds to the literature on the effectiveness of macroprudential policies by using a novel empirical setup that accounts for the simultaneous effects of these policies on the probability of financial crises and economic growth and also allows for an assessment of their net effect on the probability of a crisis. We specify an empirical framework wherein macroprudential policies exert both a direct and an indirect effect on the probability of a financial crisis through growth. In a first step, we investigate the impact of macroprudential policies on the likelihood of banking crises, specifying a dynamic logit model à la Candelon et al. (2014), where a lagged crisis index is included as regressor together with other predictors. In a second instance, we consider a bivariate vector autoregressive (VAR) model which also comprises this crisis index and output growth. This second equation (growth regression equation) is consistent with the one used in Cerrutti et al. (2017a), with the exception that regressors are lagged and not contemporaneous.[1] Once the parameter estimates of this dynamic system have been obtained via the equation-by-equation MLE approach described above, a Generalized Impulse Response Function (hereafter GIRF) analysis is conducted in order to evaluate the impact of a macroprudential policy tightening[2] on the probability of a systemic banking crisis occurring. Indeed, the GIRFs for the crisis probability and output growth measure how this bi-variate dynamic system reacts over time to an initial tightening in the macroprudential policy stance while accounting for both the direct and indirect effects of this tightening — with the former having a positive stabilizing effect which reduces the probability of occurrence of a systemic banking crisis and the latter increasing the likelihood of a future crisis occurring through its negative impact on economic growth. This setup thus allows us to quantify the net effect of macroprudential policies on financial stability by disentangling the relative contributions from the direct and indirect channels over time and by determining which effect eventually dominates.

Using a sample comprising 121 countries over the 2000-2017 period, the macroprudential policies database assembled by Cerutti et al. (2017a) and the banking crisis database of

[1]Cerutti et al. (2017a) use the Arellano and Bond (1991) GMM estimator to mitigate the endogeneity bias problem. Nevertheless, GMM based estimators suffer from a weak instrument problem in finite samples with persistent series, see e.g. Bun and Windmeijer (2010) who showed the impact of this issue in settings relevant for country-level panel data modelling. We thus adopt an MLE approach where the explanatory variables are lagged to mitigate concerns about endogeneity bias. Note that the Nickell (1981) bias affecting the fixed-effect estimates is unlikely to be a concern in our analysis given our relatively long time series dimension of 17 years.

[2]Which is modelled in our framework as a positive shock on the macroprudential policy index.

Laeven and Valencia (2018), we find that the use of macroprudential policies reduces the probability of financial crises, but that it also drags down economic growth. The negative effects of macroprudential policies on both the probability of banking crises and economic growth are economically meaningful. Our baseline estimations suggest that a one unit increase in the macroprudential policy index (MPI) lowers the conditional probability of a systemic banking crisis by 7.29 percentage points on average when we use a sample that includes both advanced and emerging market economies. A one unit rise in the MPI also lowers economic growth by 23.7 basis points. Splitting the sample into advanced and emerging market economies, we find that the effects of macroprudential policies on both growth and banking crises are most significant in emerging market economies. Turning to the net effect of macroprudential policies on the probability of financial crises, a Generalized Impulse Response Function analysis (GIRF) of a dynamic system composed of output growth and the probability of crisis reveals that the direct stabilizing effect of macroprudential policies dominates the indirect destabilizing effect.

By examining the simultaneous impact of macroprudential policies on the probability of systemic banking crises and economic growth, this paper contributes to two strands of literature. First, we add to the existing evidence documenting the benefits of macroprudential policies to financial stability. Much of the empirical literature concerned with the effectiveness of macroprudential policy has focused on assessing the impact of macroprudential tools on intermediate target financial variables, such as credit growth, credit-to-GDP, non-performing loans, and real estate asset prices (Lim et al., 2011; Cerutti et al., 2017a and 2017b; Jimenez et al., 2017; Akinci and Olmstead-Rumsey, 2018). The direct effect of macroprudential policies on the incidence of financial crises has been rarely investigated. While instructive, the effect of macroprudential policies on the behavior of intermediate financial variables is not necessarily a good indicator of their effectiveness in taming financial crises. In particular, if taken one by one, the behavior of financial variables is not necessarily an adequate indicator of the degree of financial stability or the likelihood of a crisis. In this paper, we take a different approach and contribute to the literature on macroprudential policy effectiveness by examining the direct link between the use of macroprudential tools and the incidence of financial crises. To our knowledge, this is the first paper that attempts to empirically assess the effect of various macroprudential tools on the probability of financial crisis.

Second, our study contributes to a growing body of literature that examines the real economy effects of macroprudential policy. In particular, several recent studies investigate the potential of macroprudential policies to weaken economic growth (Sanchez and Rohn, 2016; Kim and Mehrotra, 2017; Boar et al., 2017; Richter et al., 2019). Their results point to a damaging effect of macroprudential policies on economic growth. However, this literature stops short of exploring the potential effect of weaker economic performance on financial stability. We take this literature a step forward by assessing not only the impact of macroprudential policies on economic growth in a novel empirical setting, but also by considering the feedback from weaker economic growth to the probability of banking crises.

Finally, the empirical methodology is novel, borrowing to the recent literature on mixed-measurement dynamic models (Creal et al. 2014) as the dynamic system under investigation simultaneously includes a dichotomous (crisis probability) and a continuous (growth) equation. Such an approach is therefore unique and well tailored to evaluate the direct/indirect impacts of macroprudential policies on economic growth and financial stability.

The remainder of the paper is organized as follows. Section 2 provides a literature review. Section 3 presents the methodology. Results are presented in Section 4. Section 5 provides robustness analyses. Finally, Section 6 concludes.

2. Literature review

The global financial crisis (GFC) has demonstrated the importance of adopting regulatory frameworks aimed at addressing systemic financial risks. Besides traditional microprudential regulation aimed at ensuring the safety and soundness of individual institutions, such frameworks ought to encompass macroprudential policies whose goal is to enhance the resilience of the financial system as a whole. In fact, there is now a widespread recognition that, while necessary, the microprudential approach to regulation is insufficient in ensuring financial stability and preventing the incidence of costly financial crises. Increasingly more emphasis is thus put, in both academic and policy-making circles, on the use of macroprudential regulation to reduce vulnerabilities and risks in the financial system (e.g., Bank of England, 2009; Claessens et al. 2011; Hanson et al. 2011; IMF, 2013a; 2013b; IMF, 2014; ESRB, 2014; FSB, 2014; Claessens and Kodres; 2015; Freixas et al. 2015).

Despite the fact that many macroprudential tools have been in use for a long time, especially in emerging market economies (see Borio and Shim, 2007; McCauley, 2009; Lim et al. 2011), the use of such tools has intensified in the wake of the GFC in both emerging market economies (EMEs) and advanced economies (AEs). For instance, Cerutti et al. (2017a) document that macroprudential policies are used more frequently in EMEs whereas Akinci and Olmstead-Rumsey (2018) report that macroprudential policies have been used more actively after the GFC in both AEs and EMEs. Focusing on European Union countries over the period 1995-2014, Budnik and Kleibl (2018) document a widespread use of macroprudential regulations in this region and that "*the use of countercyclical policy tools has already been relatively common prior to the financial crisis*" (p.21).

Although countries across the globe have increased their use of macroprudential policies in recent years, there is no clear-cut evidence on the effectiveness of such policies in guaranteeing financial stability. Various countries, regions, periods and empirical designs have been used in studies seeking to gauge the extent to which macroprudential policy tools are successful in mitigating systemic financial risks. For instance, using IMF survey data, Lim et al. (2011) examine the effect of several macroprudential instruments on the procyclicality of credit growth.[3] Their findings suggest that these tools are effective in reducing the procyclicality of credit growth by lowering the correlation between credit growth and GDP growth.

[3]The instruments considered include caps on loan-to-value ratios, debt-to-income ratios, ceilings on credit growth, reserve requirements, time-varying/dynamic provisioning, countercyclical/time-varying capital requirements, and caps on foreign currency lending,

More recently, Cerutti et al. (2017a) examine the effectiveness of 12 macroprudential instruments in a sample of 119 countries over the 2000-2013 period. Their results point to the effectiveness of several macroprudential instruments in curbing the growth of real credit and house prices. Bruno et al. (2017) investigate the effectiveness of macroprudential and capital flow management policies in a sample of 12 Asia-Pacific countries. Their findings suggest that macroprudential policies are more successful in taming credit growth when they complement monetary policy by reinforcing monetary tightening, than when they act in opposite directions. Fondoglu (2017) assesses the effectiveness of six macroprudential policy tools in eighteen EMEs and finds that borrower-based measures are particularly successful in lowering credit growth. Akinci and Olmstead-Rumsey (2018) focus on the use of seven categories of macroprudential tools in a sample of 57 AEs and EMEs covering the period from 2000:Q1 to 2013:Q4, with episodes of tightening and loosening recorded separately. Their empirical analysis suggests that the tightening of macroprudential policies has a significant negative effect on both bank credit and house price growth. Furthermore, their results indicate that policies targeted at credit growth in certain sectors are more effective.

Another set of cross-country studies focuses on the potential role of macroprudential policy instruments in limiting house credit and the ensuing house price bubbles. In their review of country experiences with various policy options intended to contain boom-bust cycles in real estate asset markets, Crowe et al. (2013) suggest that loan-to-value (LTV) and debt-to-income (DTI) caps can be effective in reducing house price appreciation rates. Analyzing real estate booms in a sample of more than fifty countries, Cerutti et al. (2017b) find that higher LTV ratios are associated with excessively rapid house-price and credit growth during booms and suggest that LTV limits can be a well-targeted tool for limiting real-estate price fluctuations. Zhang and Zoli (2016) investigate the use of macroprudential policy instruments in 13 Asian economies and 33 economies in other regions for the period 2000-2013 and find that housing-related macroprudential measures, especially LTV caps, have been used quite extensively in Asian countries. Their analysis also suggests that housing-related macroprudential instruments — particularly LTV caps and housing tax measures — have contributed to the curbing of housing price growth, credit growth, and bank leverage in Asia. Using a sample of 57 countries, Kuttner and Shim (2016) find that housing credit growth is significantly affected by changes in the maximum DSTI ratio, the maximum LTV ratio, limits on banks' exposure to the housing sector and housing-related taxes.

Besides studies using aggregate country-level data, macroprudential policy effectiveness has also been the subject of several studies using bank- and firm-level data. Analyzing the Korean experience with macroprudential policy, Igan and Kang (2011) find that LTV and DTI limits are associated with a decline in house price appreciation and transaction activity. Claessens et al. (2013)'s analysis based on bank balance sheet data in 48 countries over 2000-2010 suggests that macroprudential policy measures aimed at borrowers — DTI and LTV caps and limits on credit growth and foreign currency lending — are effective in reducing growth in leverage, asset and noncore to core liabilities during boom times. Aiyar et al. (2014) examine micro-level evidence on the effectiveness of time-varying, bank-specific, minimum capital requirements in smoothing the credit cycle in the U.K. Their results indicate that regulated banks reduce lending in response to tighter capital requirements. Jiménez et al. (2017) use bank- and firm-level data to assess the impact of one macroprudential policy tool introduced in Spain in 2000 — dynamic provisioning — on the supply of credit during boom and bust periods. They document evidence that dynamic provisioning mitigates credit supply in good times while limiting credit crunches in bad times, which helps in smoothing downturns.

3. Methodology

To estimate the impact of macroprudential policies on the probability of financial crises, we begin with the univariate Early Warning System (EWS) model. Specifically, we use a dynamic panel logit model with fixed effects.

$$\Pr(y_{i,t} = 1) = F(\pi_{i,t}),$$
$$\pi_{i,t} = g_{i,t-1}\beta_1 + x_{i,t-1}^{\top}\beta_2 + MPI_{i,t-1}\beta_3 + \delta\pi_{i,t-1} + \eta_i, \tag{1}$$

where i is the country index and $\Pr(y_{i,t} = 1)$ is the conditional probability of observing a systemic banking crisis at time t in country i given the information at one's disposal at time $t-1$. This conditional probability relates to the crisis index, $\pi_{i,t}$, through the logistic mapping $F(\cdot)$. The dynamics of the crisis index depend on the lagged real GDP growth ($g_{i,t-1}$), the lagged macroprudential policy index ($MPI_{i,t-1}$), the lagged crisis index, a set of control variables (financial development, ratio of public debt to GPD, capital and trade openness) denoted by $x_{i,t-1}$ and a country fixed effect η_i capturing unobserved heterogeneity across countries. This univariate panel logit model can be extended to take into account economic growth (the second dimension considered by Cerrutti et al. 2017a), i.e. the growth regression. We thus extend our model to include the following equation:

$$g_{i,t} = g_{i,t-1}\Phi_1 + x_{i,t-1}^{\top}\Phi_2 + MPI_{i,t-1}\Phi_3 + \Psi\pi_{i,t-1} + c_i + \nu_{i,t}, \tag{2}$$

where c_i is a country fixed effect capturing unobserved heterogeneity across countries and $\nu_{i,t}$ is an error term. The system composed of equations (1) and (2) is estimated using the Maximum Likelihood Estimator (MLE). Our second equation resembles the model proposed by Cerrutti et al. (2017a). However, there are two fundamental differences: (i) the probability of crisis as well as growth are weakly exogenous rather than strictly exogenous and (ii) the crisis variable does not enter as a binary variable as in Cerrutti et al. (2017a), but instead as a continuous variable indicating the probability of occurrence of a crisis. This approach allows us to analyse the dynamic interactions between financial crises probability and output growth within the GIRF framework without being subject to the nonlinear "jump-like" effects that would arise from using a crisis binary variable instead.

We expect the activation of macroprudential instruments to reduce the probability of a financial crisis occurring. The MPI is thus expected to load negatively and be statistically significant in equation (1). Additionally, by constraining credit growth, the activation of macroprudential instruments should have a contractionary effect on economic activity, causing a slowdown in economic growth. In other words, we expect the MPI to have a negative and significant coefficient estimate in the growth equation (eq. 2). We subsequently use a GIRF analysis to measure the overall net effect of these two opposing effects of macroprudential policies on the probability of crises.

4. Empirical results

4.1. Data

The data we use covers the period 2000–2017 at a yearly frequency. For the banking crisis dummy, we use the Laeven and Valencia (2018) database, which records systemic financial crises across a large sample of countries. We obtain data on year-on-year growth in real GDP from the IMF World Economic Outlook (WEO) database. To extract the MPI variable, we use Cerrutti et al. (2017a)'s database covering 134 countries (33 of which are advanced economies, 63 emerging and 38 developing ones). This database relies on the IMF survey on Global Macroprudential Policy Instruments (GMPI), which provided responses to more than 100 detailed questions on the use of 17 key macroprudential policy tools. To build their composite index, Cerrutti et al. (2017a) use 12 macroprudential instruments, namely General Countercyclical Capital Buffer/Requirement (CTC); Leverage Ratio for banks (LEV); Time-Varying/Dynamic Loan-Loss Provisioning (DP); Loan-to-Value Ratio (LTV); Debt-to-Income Ratio (DTI); Limits on Domestic Currency Loans (CG); Limits on Foreign Currency Loans (FC); Reserve Requirement Ratios (RR); Levy/Tax on Financial Institutions (TAX); Capital Surcharges on SIFIs (SIFI); Limits on Interbank Exposures (INTER); and Concentration Limits (CONC). Cerrutti et al. (2017a) also define LTV CAP as the subset of LTV measures used as a strict cap on new loans, as opposed to a loose guideline or merely an announcement of risk weights; and RRREV as the subset of RR measures that impose a specific wedge on foreign currency deposits or are adjusted countercyclically. Similar to Cerrutti et al. (2017a) we also consider subsets of macroprudential instruments, separating those affecting borrowers' leverage and financial positions (LTV CAP and DTI) and those aimed at financial institutions' assets or liabilities (DP, CTC, LEV, SIFI, INTER, CONC, FC, RR REV, CG, and TAX). We also include an indicator of financial development, which we extract from Svirydzenka (2016). Ben Naceur et al. (2019) show that excessive financial development can lead to banking crises. They also show that taken individually, the various dimensions of financial development (access, depth and efficiency) are more informative than credit growth (Cerrutti et al., 2017a). Following Richter et al. (2019), and to account for external factors, we also include a country's volume of trade (in percentage of GDP) and the Chinn-Ito capital account index. Finally, data on public debt to GDP ratios is obtained from Mbaye et al. (2018) global debt database.

Table 1: Variable definitions and sources

Variable	Definition	Source
Dependent variables		
GDP growth ($g_{i,t}$)	Year on year real GDP growth (%)	IMF World Economic Outlook (WEO) database. We use GDP at constant prices.
Crisis ($y_{i,t}$)	Systemic banking crisis dummy	Laeven and Valencia (2018) crises database.
Independent variables		
FD	Financial development index	Svirydzenka (2016) aggregate index of financial development (access, depth and efficiency).
Debt-to-GDP	Public debt to GDP (%)	Mbaye et al. (2018) global debt database. We use general government debt when available and central government debt otherwise.
MPI	Macroprudential policy index (from 0 to 12)	Cerutti et al. (2017a) database. We also consider the borrower sub-index (LTV cap and DTI) and the financial sub-index (see text for details).
KA	Capital account openness index	Chinn and Ito (2006) normalized (from 0 to 1) index of capital account openness.
Trade-to-GDP	Sum of imports and exports to GDP (%)	World Bank World Development Indicators (WDI) database.

Table 2: Descriptive statistics of the data.

	Mean	Min	1st quartile	Median	3rd quartile	Max	Std. dev.	Obs.
Panel A: All countries								
GDP growth	4.06	-20.49	2.02	3.99	6.18	34.47	3.81	2178
Crisis	0.06	0.00	0.00	0.00	0.00	1.00	0.24	2178
FD	0.35	0.00	0.16	0.29	0.51	1.00	0.24	2178
Debt-to-GDP	54.33	3.09	30.08	45.07	69.59	260.96	36.13	2178
MPI	2.00	0.00	1.00	2.00	3.00	10.00	1.77	2178
MPI-Bor	0.33	0.00	0.00	0.00	0.00	2.00	0.63	2178
MPI-Fin	1.66	0.00	1.00	1.00	3.00	8.00	1.45	2178
KA	0.59	0.00	0.17	0.70	1.00	1.00	0.36	2178
Trade-to-GDP	83.36	0.17	54.98	75.09	103.41	437.33	44.66	2178
Panel B: Advanced Economies								
GDP growth	2.41	-14.81	1.14	2.45	3.93	25.01	3.33	576
Crisis	0.14	0.00	0.00	0.00	0.00	1.00	0.35	576
FD	0.66	0.20	0.56	0.70	0.81	1.00	0.19	576
Debt-to-GDP	63.65	3.66	36.31	53.86	85.29	237.65	40.68	576
MPI	1.86	0.00	1.00	2.00	3.00	6.00	1.53	576
MPI-Bor	0.45	0.00	0.00	0.00	1.00	2.00	0.68	576
MPI-Fin	1.41	0.00	0.00	1.00	2.00	6.00	1.25	576
KA	0.92	0.17	1.00	1.00	1.00	1.00	0.18	576
Trade-to-GDP	98.60	19.80	60.22	83.03	122.23	437.33	63.24	576
Panel C: Emerging Market Economies								
GDP growth	4.33	-15.10	2.39	4.23	6.26	34.47	3.80	1044
Crisis	0.04	0.00	0.00	0.00	0.00	1.00	0.19	1044
FD	0.31	0.06	0.21	0.29	0.39	0.73	0.13	1044
Debt-to-GDP	47.00	3.09	24.66	40.00	62.28	183.07	31.00	1044
MPI	2.46	0.00	1.00	2.00	3.00	10.00	1.89	1044
MPI-Bor	0.39	0.00	0.00	0.00	1.00	2.00	0.69	1044
MPI-Fin	2.06	0.00	1.00	2.00	3.00	8.00	1.51	1044
KA	0.52	0.00	0.17	0.45	0.76	1.00	0.33	1044
Trade-to-GDP	80.03	20.72	54.77	75.92	100.33	220.41	33.92	1044
Panel D:Low-Income Developing Countries								
GDP growth	5.26	-20.49	3.46	5.61	7.13	26.42	3.71	558
Crisis	0.02	0.00	0.00	0.00	0.00	1.00	0.13	558
FD	0.13	0.00	0.09	0.12	0.16	0.32	0.05	558
Debt-to-GDP	58.45	11.81	33.34	47.27	72.23	260.96	37.25	558
MPI	1.27	0.00	0.00	1.00	2.00	7.00	1.45	558
MPI-Bor	0.10	0.00	0.00	0.00	0.00	2.00	0.34	558
MPI-Fin	1.17	0.00	0.00	1.00	2.00	6.00	1.30	558
KA	0.38	0.00	0.17	0.17	0.70	1.00	0.35	558
Trade-to-GDP	73.85	0.17	48.65	67.52	95.08	206.77	34.02	558

Note: The table presents summary statistics for all observations over the period 2000-2017. Panel A presents the aggregate statistics for the 121 countries in our dataset. Panels B to D present the same summary statistics for the different country categories (respectively advanced economies, emerging market economies and low-income developing countries).

Table 2 reports descriptive statistics for all the variables used in the empirical analysis. Panel A presents these statistics for the full sample while Panels B, C and D present summary statistics for the subsamples of advanced economies, emerging market economies and low-income developing countries respectively. Table 2 suggests that systemic banking crises are most frequent in advanced economies (mean Crisis: 0.14, panel B) compared to emerging market economies (mean Crisis: 0.04) and low-income developing countries (mean Crisis: 0.02). Consistent with the findings of prior literature, on average, the use of macroprudential policy instruments is more prevalent in emerging market economies (mean MPI: 2.46) relative to both advanced economies (mean MPI: 1.86) and low-income countries (mean MPI: 1.27). We also find that borrower-based instruments are on average more used in advanced economies (mean MPI-Bor: 0.45) than in emerging market economies (mean MPI-Bor: 0.39).[4] By contrast, the use of lender-based instruments is on average more prevalent in emerging market economies (mean MPI-Fin: 2.06) than in advanced economies (mean MPI-Fin: 1.41). Economic growth rates are higher in emerging market economies and low-income developing countries compared to advanced economies.

4.2. Macroprudential policies and banking crises

Firstly, the probability of a banking crisis is estimated in a dynamic framework following Kauppi and Saikkonen (2008) and Candelon et al. (2012). Our specification captures the persistent nature of banking crises through the inclusion of the lagged crisis index.[5] The estimation of the dynamic panel logit model specified in equation (1) is carried out by Maximum Likelihood using the iterative pseudo-demeaning algorithm developed by Stammann et al. (2016). We correct for the bias in the estimated parameters — due to the incidental

[4]Note that the MPI-Bor measure is composed of only two policy instruments which results in a distribution of the policy index exhibiting a large probability mass at zero for both AEs and EMEs. This is confirmed by the reported median value of the MPI-Bor which is zero in both cases. Therefore, comparisons between the two subsamples based on the mean of the MPI-Bor might be misleading.

[5]Other types of specifications, i.e. with a lagged binary crisis variable or with both the binary variable and the index have been tested. The information criteria favor our main specification and we do not report the other ones for the sake of space. They are available from the authors upon request.

parameter problem — using the analytical bias correction derived in Hahn and Kuersteiner (2011) for general dynamic nonlinear models with fixed effects. Note that countries who did not experience at least one systemic banking crisis over our sample period do not contribute to the likelihood and are thus discarded prior to estimation. Table 3 reports the estimates obtained for all the countries as well as for the subsamples of advanced and emerging market economies.[6]

The results in Table 3 show that our main variables of interest load statistically significant and with the expected sign. Specifically, the coefficient estimate on the MPI is negative and significant, suggesting that the activation of macroprudential policies is conducive to a significant decrease in the probability of a banking crisis occurring. To assess the economic relevance of macroprudential regulation in mitigating the probability of occurrence of a systemic banking crisis, we calculate the Average Partial Effect (APE) of a one unit change in the MPI on the probability of observing a systemic banking crisis conditional on the observed values of the covariates and on the estimated parameters reported in Table 3.[7] Based on this approach, we observe that macroprudential policies also have an economically meaningful impact as a one unit change in the MPI lowers the conditional probability of a systemic banking crisis by 7.29 percentage points on average for the case in which all the countries are pooled together. The average reductions in the banking crisis probability for advanced economies and emerging market economies are respectively of 5.21 and 8.83 percentage points. Likewise, the coefficient estimate on GDP Growth is negative and significant, implying that countries with higher economic growth rates face a lower probability of banking crises. The economic effect of real economic growth on the conditional probability of a systemic banking crisis oc-

[6] As the number of countries having experienced banking crises in low income countries is very small (N = 3), the maximization algorithm does not converge and the results are thus not reported.

[7] Note that, in the case of a continuous variable $\boldsymbol{x}_k$, the Average Partial Effect (APE) is calculated as follows:

$$APE_k = \hat{\beta}_k (NT)^{-1} \sum_{t=1}^{T} \sum_{n=1}^{N} f(\hat{\pi}_{n,t}),$$

where $f(\cdot)$ is the logistic density function.

Table 3: Macroprudential Policies and Banking Crises

	All	AE	EME
MPI	−0.606***	−0.411	−0.899***
	(0.175)	(0.292)	(0.286)
GDP Growth	−0.192***	−0.283***	−0.18***
	(0.039)	(0.065)	(0.059)
Crisis Index	0.56***	0.534***	0.2
	(0.113)	(0.143)	(0.185)
FD	12.574***	11.884***	18.289**
	(3.308)	(4.06)	(7.205)
Debt-to-GDP	−0.009	−0.026**	0.013
	(0.007)	(0.012)	(0.015)
KA	−1.439	−2.058	−0.779
	(1.109)	(1.595)	(1.725)
Trade-to-GDP	0.013	0.024	0.045
	(0.012)	(0.021)	(0.028)
AIC	532.396	321.042	170.112
BIC	727.937	429.619	238.075
Pseudo R^2	0.28	0.29	0.328
# Effect. Obs.	629	357	221
# Observations	2057	544	986
Country FE	Y	Y	Y
Year FE	N	N	N

Note: ML estimation is carried out using the iterative pseudo-demeaning algorithm of Stammann et al. (2016). The dependent variable is the systemic banking crisis binary. All regressors are considered as lagged values. We report bias-corrected estimates using the analytical bias correction of Hahn & Kuersteiner (2011). Standard errors clustered at the country level are reported between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We report the Akaike (AIC) and Bayesian (BIC) information criteria together with the McFadden pseudo R squared for each specification. The estimation period is 2001-2017 and we consider one-way country fixed effects. The total and effective number of observations considered in the estimation are also reported.

curring, measured by the APE, is also sizeable. A one percentage point increase in real GDP growth rate leads to a decline of the crisis probability by 2.31 percentage points when all the countries are pooled together and by respectively 3.58 percentage points and 1.78 percentage points for AEs and EMEs. In sum, better economic performance and active macroprudential policies are important ingredients for a more stable banking sector — i.e. one that is less prone to crises. A closer look at the results reveals however that the coefficient estimate on the MPI is negative and statistically significant only in the EME subsample; while it is negative, it is not statistically significant in the AE subsample. This result is consistent with Cerutti et al. (2017a, p.211) who "*find that the statistically significant negative relation of MPI with credit growth is strongest for developing and emerging markets, and much less so for advanced economies*". The authors advance two possible reasons behind such finding: (i) the greater historical reliance of emerging markets on macroprudential policies relative to advanced countries and (ii) the fact that advanced economies tend to have more developed financial systems that offer alternative sources of finance and make regulation circumvention easier. Combined with those of Cerutti et al. (2017a), our findings can be interpreted as evidence that macroprudential policies have been used effectively by emerging market countries to tame credit growth, which eventually led to a lower incidence of banking crises. The upshot of this analysis is that macroprudential policies have a meaningful positive impact on financial stability in emerging market economies while their role in fostering financial stability in advanced economies is less certain. Economic growth is however fundamental for financial stability in both groups of countries. Also, and as expected, the coefficient on the autoregressive index (Crisis Index) is lower than 1, meaning that there is no absorbing state of crisis. This coefficient is significantly different from 0 indicating that a static model would be misspecified. Consistent with the results reported by Ben Naceur et al. (2019), we find that greater financial development is conducive to a higher probability of a banking crisis; the coefficient estimate on the FD index is positive and significant in the full sample as well as across the two subsamples of AEs and EMEs. The other macroeconomic variables

(Debt-to-GDP, KA and Trade openness) do not seem to matter for the likelihood of banking crises. Out of our set of explanatory variables, macroprudential policies, economic growth and financial development appear to be good predictors of banking crises.

4.3. Macroprudential policies, economic growth and banking crises

Secondly, we estimate the growth regression equation including the same explanatory variables as in the banking crisis equation estimated previously. It is worth clarifying that the probability of crisis can be introduced via the latent index estimated in the first equation; Crisis Index is thus a continuous variable taking values on the real line. Such an approach differs from Cerrutti et al. (2017a) who include information about past crises as a binary variable. Results of the estimation are reported in Table 4.

Consistent with our expectations, the coefficient estimate on the MPI turns out to be negative and statistically significant, suggesting that the activation of macroprudential policies depresses economic growth. The effect of the MPI on economic growth is also economically meaningful; in the full sample, a one unit increase in the MPI lowers economic growth by 23.7 basis points. In the EME subsample, a one unit increase in the MPI depresses economic growth by 31.8 basis points. This result is consistent with Richter et al. (2019)'s findings that macroprudential measures, specifically changes in the maximum loan-to-value ratio, have a significant negative effect on output. It should, however, be noticed that this result is mainly driven by emerging market countries as the MPI does not load statistically significant in the AE subsample. For this latter group of countries, the activation of macroprudential policies does not seem to bear a cost in terms of economic growth. We can thus conclude from these preliminary evidence that the indirect channel of transmission of macroprudential policies to the likelihood of occurrence of a systemic banking crisis, through their impact on economic growth, is operative in emerging market economies but rather muted in advanced economies.

The coefficient estimate for the banking crisis index is negative and significant in both the full sample and the AE subsample. This indicates that banking crises, especially in advanced economies, put a drag on economic growth. This result is in line with the evi-

Table 4: Macroprudential Policies and Real GDP growth

	All	AE	EM
MPI	-0.237^{***}	-0.037	-0.318^{***}
	(0.079)	(0.132)	(0.122)
GDP Growth	0.306^{***}	0.368^{***}	0.372^{***}
	(0.022)	(0.044)	(0.031)
Crisis Index	-0.204^{**}	-0.202^{*}	-0.147
	(0.092)	(0.117)	(0.150)
FD	-6.320^{***}	-6.863^{**}	-5.510^{**}
	(1.792)	(2.977)	(2.446)
Debt-to-GDP	0.011^{***}	0.023^{***}	0.028^{***}
	(0.004)	(0.009)	(0.008)
KA	-0.541	0.969	-0.766
	(0.627)	(1.379)	(0.778)
Trade-to-GDP	0.002	-0.031^{***}	0.021^{**}
	(0.005)	(0.011)	(0.009)
Country FE	Y	Y	Y
Year FE	N	N	N
Observations	2,057	544	986
R^2	0.361	0.299	0.347
Adjusted R^2	0.319	0.247	0.301

Note: Estimation is carried out by OLS with one-way country fixed effect dummies. The estimation period is 2001-2017. The dependent variable is the year-on-year real GDP growth rate. All regressors are considered as lagged values. We report estimates with Standard errors clustered at the country level between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We report both the R-squared and adjusted R-squared for each specification.

dence documented by Candelon et al. (2016) and Candelon et al. (2019) that advanced economies' economic performance is adversely affected by banking crises whereas low income and emerging market economies' economic performance mainly suffers from currency crises.

Contrary to the banking crisis equation results, some macroeconomic control variables turn out to be statistically significant. Debt-to-GDP loads positively indicating that countries with higher public debt relative to GDP enjoy higher economic growth. Trade openness seems to bear contradicting effects on economic growth depending on whether the economy is an advanced one or an emerging one. More trade openness spurs economic growth in emerging markets while it drags it down in advanced economies.

We also carried out the estimation of Equation (2) using the GMM estimator proposed in Arellano and Bond (1991) to investigate the sensitivity of our main conclusions to endogeneity issues.[8] The results are reported in Table B.10 of Appendix B together with a short summary of the main findings. The analysis indicates that our main conclusions on the significance of the indirect channel of transmission of macroprudential policies are robust to the estimation approach adopted. The analysis in Section 4.4 on the direct and indirect effects of macroprudential regulation based on the MLE estimates might understate the indirect channel of transmission. This could result in more conservative estimates of the response in economic activity to a tightening in the macroprudential stance. However, our results would remain qualitatively robust to different estimates on the strength of the indirect channel of transmission for macroprudential policies.

4.4. Direct versus indirect effects: a multivariate analysis

Our empirical analyses in the two previous subsections suggest that macroprudential policies can have an ambiguous impact on the incidence of banking crises. On the one hand, by mitigating vulnerabilities and risks in the banking sector, macroprudential policies decrease the probability of occurrence of a crisis. We call this the direct effect of macroprudential policies

[8]This analysis is nevertheless subject to the caveat mentioned in the introduction about the weak instrument problem of GMM estimators in finite samples. See footnote 1 for more details.

on banking crises (equation 1 effect). On the other hand, our estimation results for equation (2) reveal that macroprudential policies tend to depress economic growth, which according to the parameter estimates in equation (1) would raise the probability of financial crises occurring; i.e. the lower the economic growth and the higher is the probability of a banking crisis. We call this effect of macroprudential policies on the probability of crises, which works through economic growth, the indirect effect. It is thus interesting to further investigate this issue to uncover the net effect of macroprudential policies on financial stability. In particular, it is of interest to understand whether the stabilizing direct effect of macroprudential policies dominates or whether it is the destabilizing indirect effect that dominates. If the direct effect dominates, one should observe a positive net effect of macroprudential policies on financial stability — a lower probability of crisis. Inversely, if the indirect effect dominates, one should see a negative net effect of macroprudential policies on financial stability — a higher probability of crisis. To this end, we propose a Generalized Impulse Response Function (GIRF) analysis that disentangles these two effects and uncovers the time horizon within which the domination of one effect over the other takes place.[9]

We report the impact of a macroprudential policy tightening, measured by a positive variation in the MPI, on the conditional probability of a systemic banking crisis and on real GDP growth. In particular, a variation by one unit of the MPI corresponds to the activation of an additional macroprudential policy instrument. Examples of such activations include the introduction of a cap on loan-to-value ratios or the introduction of a capital surcharge for SIFIs. In this respect, our definition of a macroprudential policy tightening (i.e. of a positive shock) involves a substantial change in policy stance and has to be differentiated from other milder types of policy interventions such as further recalibrations/adjustments of existing policy instruments.[10] In order to give further insights on the magnitude of the

[9] Appendix C describes in details the method used to obtain it at an horizon h.

[10] To illustrate, let us consider two examples of LTV policy actions conducted in South Korea: the introduction on September 9, 2002 of a maximum LTV ratio of 60% on real estate loans would be considered as a macroprudential tightening (MPI increasing by 1). On the contrary, the increase in the LTV ratio from 60 to 70% for long maturity loans introduced on March 24, 2004 would not be considered as a tightening (MPI

macroprudential policy tightenings, we refer to the descriptive statistics of the macroprudential policy index reported in Table 2. We see that in the case of advanced economies, the MPI can take values ranging from 0 to 6 with a mean of 1.86 and a standard deviation of 1.53. For emerging economies, we observe that the distribution of MPI is skewed towards higher values. Indeed, the MPI values range from 0 to 10 with a mean of 2.46 and a standard deviation of 1.89. Given these summary statistics, we will refer to the three cases corresponding to a variation of the MPI by 1, 2 and 3 (designated by *delta* in the columns of Figures 1 and 2) as respectively *small*, *moderate*, and *large* macroprudential policy tightenings. To summarize, our GIRF analysis quantifies how imposing an increase in the MPI at time $h = 0$ would conditionally impact the evolution of the crisis probability and output growth at times $h = 0, \ldots, H$ — obtained from the dynamic system defined by Equations (1) and (2) — compared to a baseline scenario where no initial tightening in the MPI took place. In addition, we quantify the uncertainty around these estimates through a Moving-Block Bootstrap (MBB) procedure which takes into account both the time series dependence in the data and the cross-sectional dimension.[11]

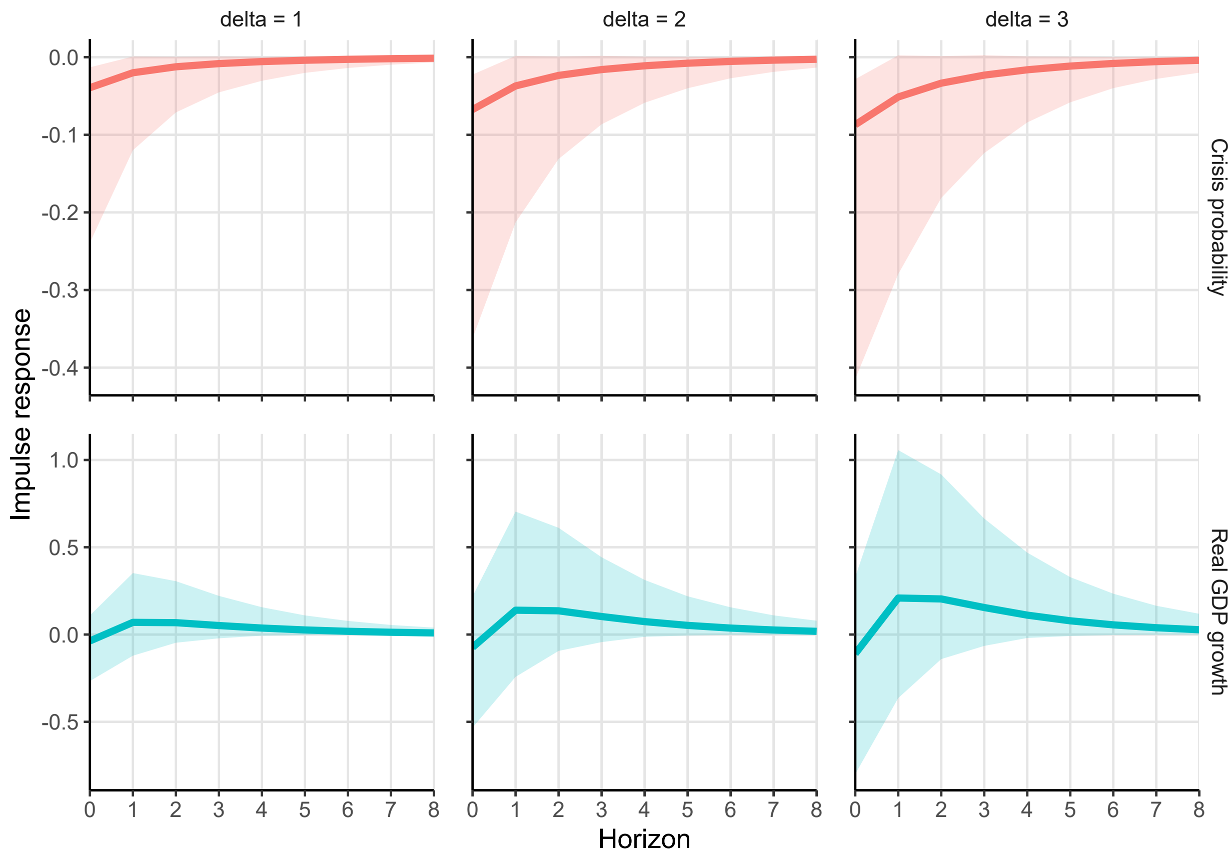
The results of our analysis are presented respectively in Figures 1 and 2 for advanced economies and emerging market economies. For advanced economies, we see in the upper panel of Figure 1 that the impact of a macroprudential policy tightening on the conditional probability of a systemic banking crisis is always negative. The crisis probability decreases just after the implementation of a MPI tightening by respectively 3.93 percentage points for a small policy tightening, by 6.74 percentage points for a moderate tightening, and by 8.7 percentage points for a large policy tightening. Furthermore, these results are rather imprecisely estimated as the 68% Confidence Interval (CI) is consistent with an effect of a macroprudential policy tightening on the conditional probability of systemic banking crisis ranging from -1.31 to -23.9 percentage points in the small tightening case, from -2.24 to -36.21 per-

does not change). Further information about the construction of the MPI index can be found in Cerutti et al. (2017a), pages 205–206.

[11] The procedure for the Moving-Block Bootstrap is detailed in Appendix D.

centage points in the moderate tightening case, and from -2.84 to -41.48 percentage points in the large tightening case. After the initial introduction of the macroprudential policy tightening, we do not find any statistically significant effect of macroprudential regulation on the probability of crisis at horizons ranging from 1 to 8 years. Moving on to the lower panel of Figure 1, we observe that there is no evidence of a significant effect of a macroprudential policy tightening on real GDP growth if we consider a 68% confidence interval for inference.

Figure 1: GIRF analysis of a macroprudential tightening: advanced economies

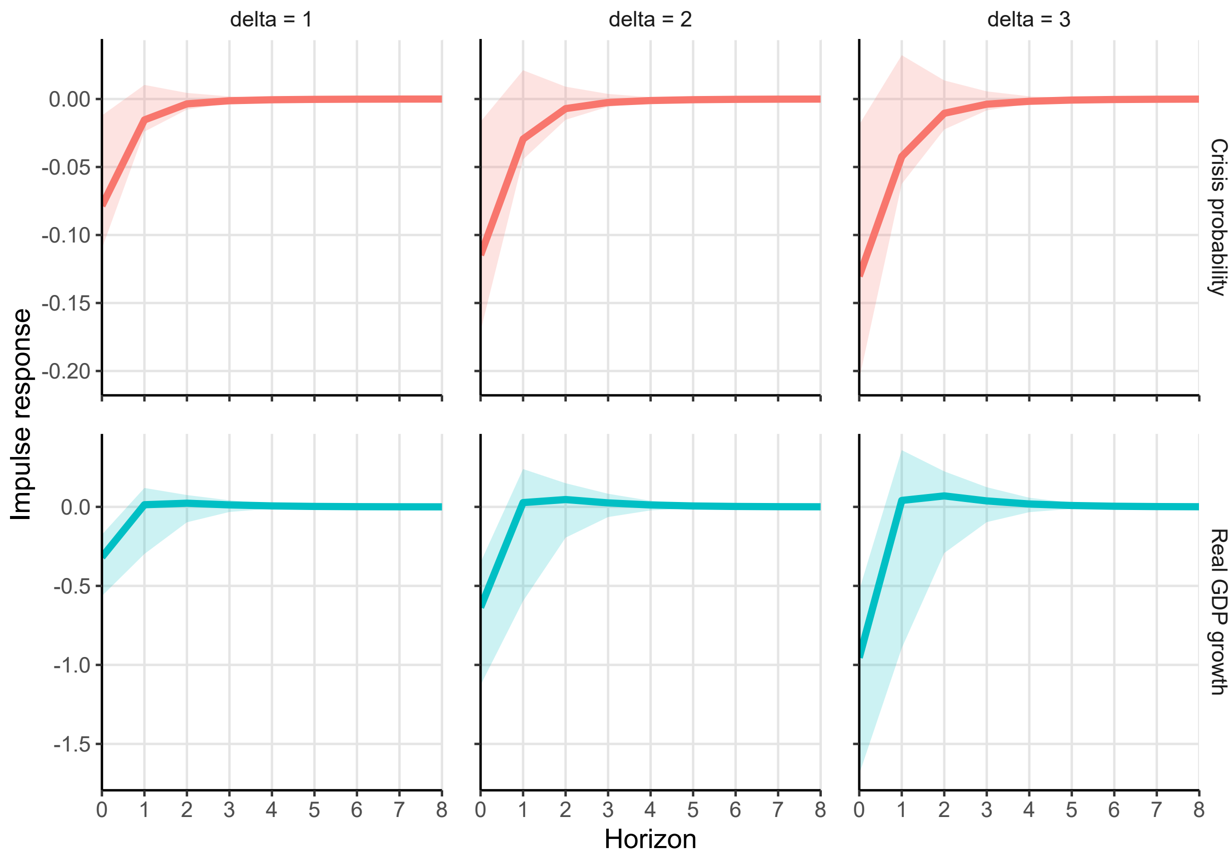


Note: Comparison of the impulse responses of the probability of a systemic banking crisis (top-panel in red) and of real GDP growth (bottom-panel in green) to different levels of macroprudential tightenings (positive MPI shocks denoted by delta). The horizon of the responses displayed ranges up to eight years after the policy tightening. The Generalized Impulse Response Functions are computed following the procedure outlined in Appendix C. The shaded areas around the response functions correspond to the 68% confidence intervals from the Moving-Block Bootstrap procedure detailed in Appendix D with 5000 bootstrap replications and a block size of 5 years.

Taken together, these results for the impact of a MPI tightening on the conditional probability of a systemic banking crisis and on real GDP growth lead us to conclude that, in the case of advanced economies, the stabilizing direct effect of macroprudential policies tend to dominate as we observe a statistically significant reduction in the probability of crisis following a macroprudential policy tightening.

We now move to the analysis of the impact of a macroprudential policy tightening in the case of emerging market economies. The results are displayed in Figure 2. We see in the upper panel that the conditional probability of a systemic banking crisis decreases significantly after the introduction of a macroprudential policy tightening. The initial estimated decrease in crisis probability is 7.85 percentage points for a small policy tightening, 11.47 percentage points for a moderate tightening, and 13.02 percentage points for a large policy tightening. We thus observe that the estimated effect of a MPI tightening is larger for emerging market economies than for advanced economies. In addition, these results are more precisely estimated than in the case of advanced economies, as indicated by the narrower 68% confidence intervals. The 68% CI is consistent with an effect of a macroprudential policy tightening on the conditional probability of systemic banking crisis ranging from -1.21 to -11 percentage points in the small tightening case, from -1.67 to -17.06 percentage points in the moderate tightening case, and from -1.88 to -20.6 percentage points in the large tightening case. Similar to our results for advanced economies, we do not find any statistically significant effect of macroprudential regulation on the probability of crisis after the initial introduction of the macroprudential policy tightening.

Figure 2: GIRF analysis of a macroprudential tightening: emerging market economies



Note: Comparison of the impulse responses of the probability of a systemic banking crisis (top-panel in red) and of real GDP growth (bottom-panel in green) to different levels of macroprudential tightenings (positive MPI shocks denoted by delta). The horizon of the responses displayed ranges up to eight years after the policy tightening. The Generalized Impulse Response Functions are computed following the procedure outlined in Appendix C. The shaded areas around the response functions correspond to the 68% confidence intervals from the Moving-Block Bootstrap procedure detailed in Appendix D with 5000 bootstrap replications and a block size of 5 years.

Turning to the impact of macroprudential regulation on real economic activity, we see in the lower panel of Figure 2 that real GDP growth experiences a contraction after the introduction of a macroprudential policy tightening. The initial decrease in real GDP growth is estimated to be 31.8 basis points for a small policy tightening, 63.6 basis points for a moderate tightening and 95.4 basis points for a large policy tightening. This initial drop in economic activity is also statistically significant as indicated by the 68% CIs which support an effect of a macroprudential tightening on real GDP growth ranging from -17.5 to -56.4 basis points in the small tightening case, from -35.1 to -112.7 basis points in the moderate tightening case, and from -52.6 to -169.1 basis points in the large tightening case. We do not find any statistically significant effect of macroprudential regulation on real economic activity for the periods following the initial introduction of the policy tightening.

We conclude that tighter macroprudential regulation in emerging market economies is improving financial stability at the cost of harming substantially economic growth at the time of the introduction of the stricter regulation. Nevertheless, the stabilizing direct effect of macroprudential policies seems also to dominate the indirect harmful effect on economic activity for emerging market economies.

In summary, our analysis suggest that despite the negative impact on growth, macroprudential policies are overall beneficial with respect to financial stability. In particular, there is a net positive effect of macroprudential policies on financial stability — reflected by a lower probability of systemic banking crises — in both advanced and emerging market economies. This positive effect on financial stability is also observed regardless of the size of the shock to macroprudential policy. Be it in advanced or in emerging market economies, a macroprudential policy tightening is thus likely to decrease the future occurrence of a banking crisis.

5. Robustness analysis

In this section, we conduct two robustness checks on the results obtained in the previous sections. First, we investigate whether the results we obtained for the impact of macroprudential policies on real GDP growth and on the probability of a systemic banking crisis change when we consider only a specific subset of macroprudential instruments. We focus on policies specifically aimed at borrowers' leverage and financial positions and on policies aimed at financial institutions' assets or liabilities. Second, we want to assess whether the Global Financial Crisis had an impact on the transmission of macroprudential regulation to systemic banking crises and real economic growth. We present the results of these two robustness checks in the following subsections.

5.1. Targeting borrowers vs. financial institutions

Following Cerutti et al. (2017a), we further divide the MPI into a subset of macroprudential instruments targeted at borrowers' leverage and financial positions[12], which we label MPI-Bor and a subset of macroprudential instruments targeted at financial institutions' assets or liabilities[13], which we label MPI-Fin.

We estimate again the dynamic panel logit specification reported in equation (1) on our full sample of countries and on the subgroups of advanced and emerging market economies. The estimation method, data filtering, and bias-correction are as described in Section 3 and we refer the reader to section 4.2 for details. The results are reported in Table 5 below. They confirm our main result that macroprudential policy tools lead to a significant decrease in the probability of systemic banking crisis. We also find that our result for the aggregate MPI is mostly driven by the effect of macroprudential instruments targeted at financial institutions' assets or liabilities — MPI-Fin. Indeed, the MPI-Bor does not contribute significantly to the

[12] This subgroup includes loan-to-value caps and debt-to-income ratios.

[13] This subgroup contains dynamic loan-loss provisioning, countercyclical capital buffers, leverage ratios for banks, capital surcharges on SIFIs, limits on interbank exposures, concentration limits, limits on domestic and foreign currency loans, reserve requirements, and tax on financial institutions.

conditional probability of a systemic banking crisis in advanced economies. MPI-Fin, on the other hand, significantly reduces the probability of a banking crisis for both advanced and emerging economies. These findings help thus in refining our analysis based on the overall MPI and improve our understanding of the particular types of macroprudential tools that lower the probability of the incidence of banking crises. While in aggregate the positive impact of macroprudential policies on financial stability is unambiguous for emerging market economies, this positive impact is mostly driven by policies directed towards financial institutions' assets or liabilities in the case of advanced economies. Our results thus complement and extend the analysis in Cerutti et al. (2017a) who document that the measures associated with the MPI-Bor and MPI-Fin significantly contribute to dampening credit growth in the full sample of countries and the subsample of EMEs but not in the subsample of AEs. We document that the measures associated with the MPI-Fin are also effective in promoting financial stability for AEs by significantly reducing the incidence of systemic banking crises.

Additionally, We observe that the economic significance of borrower- and financial-oriented macroprudential policies is even stronger than in our baseline case with the aggregate MPI. On average, a one unit change in the MPI-Bor lowers the probability of crisis by 8.91 percentage points when all countries are pooled together and by 16.21 percentage points when only EMEs are included in the sample. Likewise, a one unit change in MPI-Fin reduces the conditional probability of a banking crisis respectively by 9.38 percentage points when the sample includes all the countries, by 7.43 percentage points for the AEs subsample and by 8.91 percentage points for the EMEs subsample.

The results for the statistically significant negative effect of real GDP growth on the conditional probability of a systemic banking crisis are confirmed for both the MPI-Fin and the MPI-Bor indices. Furthermore, the economic effect of real economic growth on the conditional probability of a systemic banking crisis is comparable to our baseline results with the aggregate macroprudential policy index. We thus confirm our result that sound economic growth is fundamental for the stability of the financial sector.

Table 5: Macroprudential Policies and Banking Crises: borrower and financial indices

	All	AE	EME	All	AE	EME
MPI-Bor	-0.705^{*}	0.368	-1.556^{**}			
	(0.428)	(0.719)	(0.651)			
MPI-Fin				-0.78^{***}	-0.59^{*}	-1.094^{***}
				(0.219)	(0.343)	(0.356)
GDP Growth	-0.179^{***}	-0.282^{***}	-0.174^{***}	-0.188^{***}	-0.28^{***}	-0.176^{***}
	(0.038)	(0.065)	(0.057)	(0.04)	(0.066)	(0.06)
Crisis Index	0.532^{***}	0.492^{***}	0.04	0.553^{***}	0.532^{***}	0.2
	(0.124)	(0.139)	(0.214)	(0.114)	(0.144)	(0.194)
FD	11.061^{***}	13.465^{***}	15.464^{**}	12.406^{***}	11.929^{***}	15.602^{**}
	(3.156)	(4.174)	(6.852)	(3.243)	(4.069)	(6.422)
Debt-to-GDP	-0.009	-0.025^{**}	0.021	-0.009	-0.027^{**}	0.008
	(0.006)	(0.011)	(0.016)	(0.007)	(0.012)	(0.015)
KA	-1.212	-3.037^{*}	-1.351	-1.882^{*}	-2.341	-1.364
	(1.102)	(1.645)	(1.699)	(1.086)	(1.545)	(1.734)
Trade-to-GDP	0.01	0.009	0.066^{**}	0.013	0.029	0.043
	(0.012)	(0.017)	(0.028)	(0.012)	(0.021)	(0.028)
AIC	560.732	326.489	180.955	531.424	317.975	172.158
BIC	756.273	435.065	248.919	0.282	0.298	0.318
Pseudo R^2	0.234	0.275	0.271	0.282	0.298	0.318
# Effect. Obs.	629	357	221	629	357	221
# Observations	2057	544	986	2057	544	986
Country FE	Y	Y	Y	Y	Y	Y
Year FE	N	N	N	N	N	N

Note: ML estimation is carried out using the iterative pseudo-demeaning algorithm of Stammann et al. (2016). The dependent variable is the systemic banking crisis binary. All regressors are considered as lagged values. We report bias-corrected estimates using the analytical bias correction of Hahn & Kuersteiner (2011). Standard errors clustered at the country level are reported between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We report the Akaike (AIC) and Bayesian (BIC) information criteria together with the McFadden pseudo R squared for each specification. The estimation period is 2001-2017 and we consider one-way country fixed effects. The total and effective number of observations considered in the estimation are also reported.

We also re-estimate equation (2) for real GDP growth by considering borrower- and financial-based macroprudential policies instead of the aggregate MPI. The results are reported in Table 6. Consistent with our baseline result using a sample that includes all the countries, both borrower- and financial-based macroprudential policies have a significantly negative impact on real economic growth. A closer look at Table 6 also reveals that the MPI-Fin results are driven by emerging market countries. The economic significance of the macroprudential policies targeting borrowers and financial institutions is also stronger than in the case of the aggregate MPI. A one unit change in the MPI-Bor leads to a contraction of real GDP growth by 36.6 basis points for the sample including all countries. In the case of the MPI-Fin, a one unit increase translates in a contraction of real economic activity by respectively 28.8 basis points when all the countries are pooled together and 36.8 basis points for emerging market economies.

Finally, the statistical significance of other macroeconomic control variables such as Debt-to-GDP (positive sign) and Trade-to-GDP (negative sign for AEs and positive sign for EMEs) is consistent with our baseline estimations reported previously.

We now turn to the assessment of the impact of financial- and borrower-based macroprudential policies considered separately on the trade-offs between the direct and indirect effects of macroprudential regulation outlined in section 4.4. First, the descriptive statistics reported in Table 2 provide additional information on the distribution of the MPI-Bor and MPI-Fin variables. The MPI-Fin in advanced economies takes values ranging from 0 to 6 with a mean of 1.41 and a standard deviation of 1.25. In EMEs, its values are comprised between 0 and and 8 with a mean of 2.06 and a standard deviation of 1.51. The MPI-Bor, which is only composed of two categories of macroprudential instruments, has de facto a much narrower range of variations with a large amount of country-year observations taking a value of zero. Given that the aggregate MPI is composed of ten instruments targeted at the financial sector out of a total of twelve and that the distribution of the MPI-Fin in advanced economies and emerging market economies is comparable to the aggregate MPI, we will keep our terminology of *small*, *moderate*, and *large* policy tightenings to refer to variations of the MPI-Fin by respectively 1, 2, and 3. In the interest of space, we will focus on the analysis of the results for the MPI-Fin in the rest of this section.[14] The results for advanced and emerging market economies are presented respectively in Figures 3 and 4 below.

Table 6: Macroprudential Policies and Real GDP growth: borrower and financial indices

	All	AE	EM	All	AE	EM
MPI-Bor	-0.366^{**}	-0.191	-0.439			
	(0.186)	(0.268)	(0.282)			
MPI-Fin				-0.288^{***}	0.019	-0.368^{**}
				(0.103)	(0.183)	(0.156)
GDP Growth	0.307^{***}	0.361^{***}	0.379^{***}	0.306^{***}	0.367^{***}	0.374^{***}
	(0.022)	(0.044)	(0.031)	(0.022)	(0.044)	(0.031)
Crisis Index	-0.274^{**}	-0.231^{**}	-0.032	-0.194^{**}	-0.200^{*}	-0.106
	(0.114)	(0.115)	(0.183)	(0.091)	(0.115)	(0.159)
FD	-6.827^{***}	-6.632^{**}	-7.316^{***}	-6.512^{***}	-6.787^{**}	-5.887^{**}
	(1.765)	(2.972)	(2.297)	(1.785)	(2.974)	(2.427)
Debt-to-GDP	0.011^{***}	0.024^{***}	0.028^{***}	0.011^{***}	0.023^{**}	0.028^{***}
	(0.004)	(0.009)	(0.008)	(0.004)	(0.009)	(0.008)
KA	-0.593	0.893	-0.786	-0.545	0.886	-0.765
	(0.628)	(1.373)	(0.780)	(0.628)	(1.377)	(0.781)
Trade-to-GDP	0.001	-0.030^{***}	0.020^{**}	0.002	-0.032^{***}	0.020^{**}
	(0.005)	(0.010)	(0.009)	(0.005)	(0.011)	(0.009)
Country FE	Y	Y	Y	Y	Y	Y
Year FE	N	N	N	N	N	N
Observations	2,057	544	986	2,057	544	986
R^2	0.361	0.301	0.344	0.361	0.300	0.346
Adjusted R^2	0.319	0.249	0.298	0.319	0.247	0.300

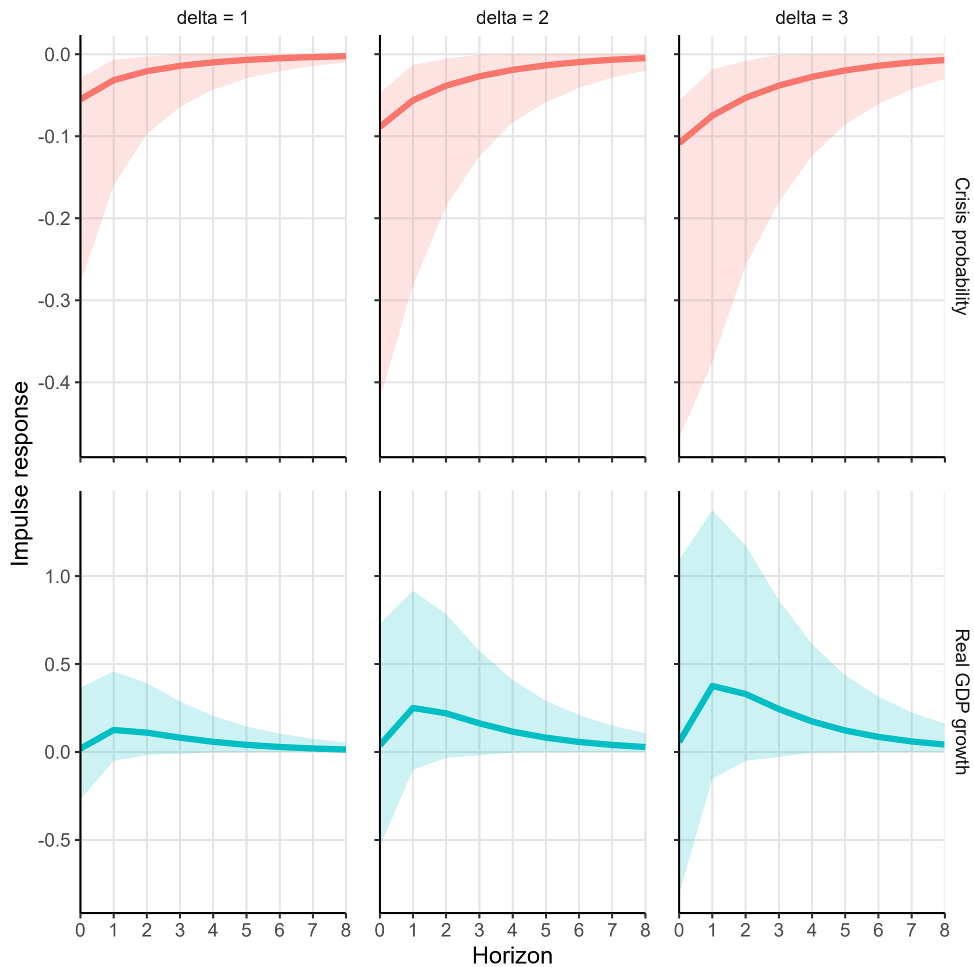
Note: Estimation is carried out by OLS with one-way country fixed effect dummies. The estimation period is 2001-2017. The dependent variable is the year-on-year real GDP growth rate. All regressors are considered as lagged values. We report estimates with Standard errors clustered at the country level between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We report both the R-squared and adjusted R-squared for each specification.

For advanced economies, we see in the upper panel of Figure 3 that the introduction of a macroprudential policy tightening targeted at the financial sector produces a reduction in the

[14]We also computed the GIRFs for the MPI-Bor. Although we didn't find any statistically significant effect of a MPI-Bor tightening on crisis probability or on real economic activity, the results are provided in Appendix E for completeness.

conditional probability of a systemic banking crisis which is quantitatively similar to the one obtained with the aggregate macroprudential policy index. The crisis probability decreases just after the implementation of a MPI-Fin tightening by respectively 5.50 percentage points for a small policy tightening, by 8.87 percentage points for a moderate tightening, and by 10.84 percentage points for a large policy tightening. Similar to our results with the aggregate MPI, these impulse responses are rather imprecisely estimated as shown by the width of the reported 68% confidence intervals. A major difference in the results for the MPI-Fin compared to the aggregate MPI is that the reduction in the conditional probability of a conditional systemic banking crisis is still statistically significant at horizons of up to two years. The crisis probability two years after the introduction of the MPI-Fin tightening is lower than the value in the case of no policy change respectively by 2.1 percentage points for a small policy tightening, by 3.83 percentage points for a moderate tightening, and by 5.31 percentage points for a large policy tightening. We do not find any evidence of a significant effect of a macroprudential policy tightening targeted at the financial sector on real GDP growth. This result is consistent with our analysis for the aggregate macroprudential policy index. Furthermore, the width of the 68% confidence intervals reported for the real GDP growth impulse responses in the lower panel of Figure 3 indicate that the results are subject to more estimation uncertainty than in the aggregate MPI case.

Figure 3: GIRF analysis of a financial macroprudential tightening: advanced economies

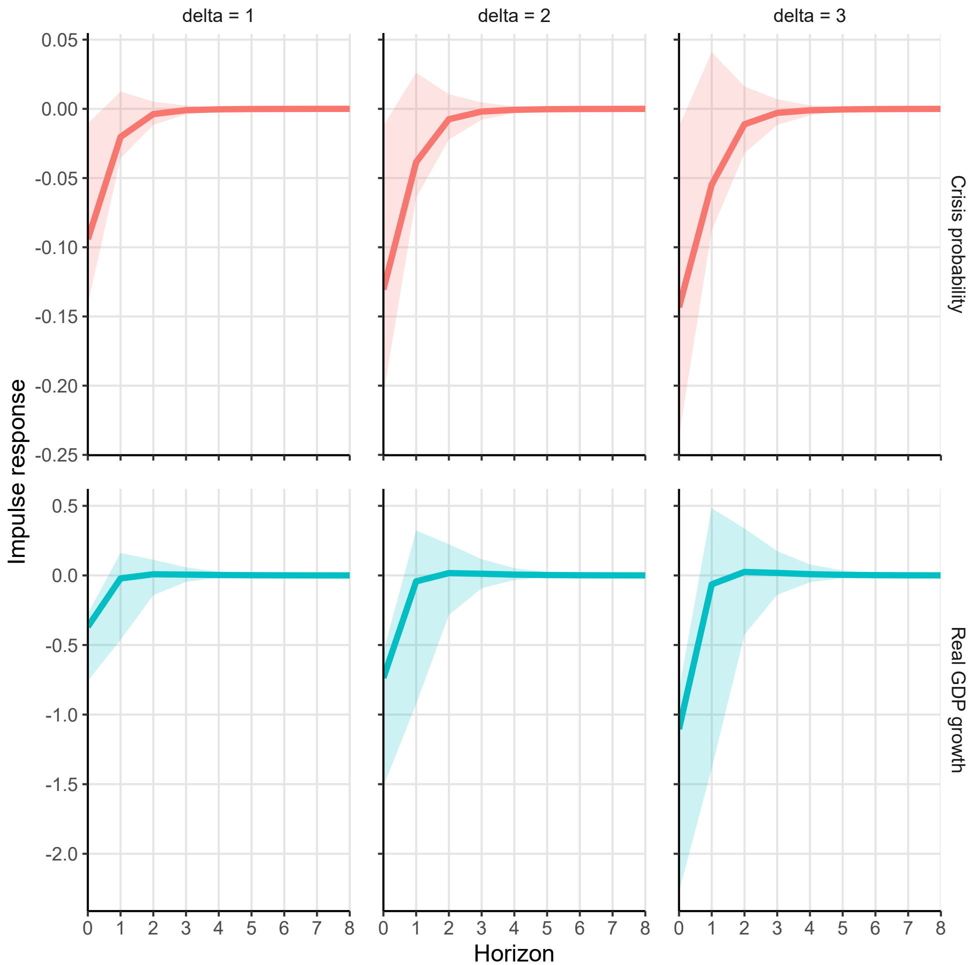


Note: Comparison of the impulse responses of the probability of a systemic banking crisis (top-panel in red) and of real GDP growth (bottom-panel in green) to different levels of financial macroprudential tightenings (positive MPI-Fin shocks denoted by delta). The horizon of the responses displayed ranges up to eight years after the policy tightening. The Generalized Impulse Response Functions are computed following the procedure outlined in Appendix C. The shaded areas around the response functions correspond to the 68% confidence intervals from the Moving-Block Bootstrap procedure detailed in Appendix D with 5000 bootstrap replications and a block size of 5 years.

We now shift our focus to the results obtained for EMEs and reported in Figure 4 below. We observe in the upper panel that the decrease in the systemic banking crisis probability after the introduction of a macroprudential policy tightening targeted at the financial sector is statistically significant and its magnitude is comparable to the results obtained in section 4.4 for the aggregate macroprudential policy index. We do not find any statistically significant effect of macroprudential regulation on the probability of crisis after the initial introduction of the policy tightening. The results for real GDP growth reported in the lower panel, on the other hand, highlight that the decrease in real economic activity induced by the introduction of a tightening in the financial macroprudential index is more pronounced than in the case of the aggregate MPI. The initial decrease in real GDP growth is estimated to be 36.8 basis points for a small policy tightening, 73.62 basis points for a moderate tightening and 110.43 basis points for a large policy tightening. This initial drop in economic activity is concentrated around more negative values as indicated by the 68% CIs which support an effect of a MPI-Fin tightening on real GDP growth ranging from -27.42 to -75.8 basis points in the small tightening case, from -54.83 to -151.59 basis points in the moderate tightening case, and from -82.25 to -227.39 basis points in the large tightening case. No further statistically significant effect is found at horizons longer than the initial introduction of the policy tightening.

In conclusion of this first robustness check, we confirm that there is a net positive effect of macroprudential policies targeted at the financial sector on financial stability in both advanced and emerging market economies. A macroprudential policy tightening targeting the financial sector decreases the future occurrence of a banking crisis, even after taking into account its indirect harmful effect on real economic activity. In the case of advanced economies, this reduction in the conditional probability of a systemic banking crisis persists at horizons of up to two years after the introduction of the policy tightening.

Figure 4: GIRF analysis of a financial macroprudential tightening: emerging market economies



Note: Comparison of the impulse responses of the probability of a systemic banking crisis (top-panel in red) and of real GDP growth (bottom-panel in green) to different levels of financial macroprudential tightenings (positive MPI-Fin shocks denoted by delta). The horizon of the responses displayed ranges up to eight years after the policy tightening. The Generalized Impulse Response Functions are computed following the procedure outlined in Appendix C. The shaded areas around the response functions correspond to the 68% confidence intervals from the Moving-Block Bootstrap procedure detailed in Appendix D with 5000 bootstrap replications and a block size of 5 years.

5.2. *The impact of the Global Financial Crisis*

As highlighted in the introduction, the Global Financial Crisis (GFC) prompted public authorities around the world to reassess their existing toolbox of macroprudential instruments and to implement new measures to reduce the future occurrence of systemic banking crises (see Claessens and Kodres (2015) for the regulatory responses to the GFC).

In this second robustness check, our aim is to test whether the changes in both the scale and composition of the macroprudential policy landscape in response to the GFC had a significant impact on the transmission of macroprudential regulation to the conditional probability of systemic banking crisis and to real economic activity. To this end, we modify equations (1) and (2) to account for a potential change after 2008 in the sensitivity of the conditional probability of systemic banking crisis and of real GDP growth to changes in the macroprudential policy index, MPI. The specification of the linear systemic banking crisis index, $\pi_{i,t}$, outlined in equation (1) is modified as follows

$$\begin{aligned}\pi_{i,t} &= g_{i,t-1}\beta_1 + x_{i,t-1}^{\top}\beta_2 + MPI_{i,t-1}\beta_3 \\ &\quad + \delta\pi_{i,t-1} + \mathbb{1}_{\{t-1\ \geq\ 2008\}}\ MPI_{i,t-1}.\beta_4 + \eta_i, \end{aligned} \tag{3}$$

where we added an interaction variable between the MPI variable and an indicator variable, $\mathbb{1}_{\{t-1\ \geq\ 2008\}}$, taking a value of 0 before 2008 and of 1 after 2008 to capture the potential change in the transmission of macroprudential policies after the GFC. Similarly, an interaction variable between the MPI and a post-GFC indicator function is added to the growth regression specified in equation (2):

$$\begin{aligned}g_{i,t} &= g_{i,t-1}\Phi_1 + x_{i,t-1}^{\top}\Phi_2 + MPI_{i,t-1}\Phi_3 \\ &\quad + \Psi\pi_{i,t-1} + \mathbb{1}_{\{t-1\ \geq\ 2008\}}\ MPI_{i,t-1}\Phi_4 + c_i + \nu_{i,t}. \end{aligned} \tag{4}$$

We estimate the dynamic panel logit model with the modified specification that we introduced in equation (3) on our full sample of countries and on the subgroups of advanced and emerging market economies. The results are reported in the right panel of Table 7 below. We also report the results of our baseline specification (equation 1) in the left panel for ease of comparison.

Our estimations reveal that the impact of the GFC on the transmission of macroprudential policy to financial stability, as measured by the conditional probability of a systemic banking crisis, differs across advanced economies and EMEs. For advanced economies, the GFC did not significantly impact the ability of macroprudential policies to reduce the conditional probability of a systemic banking crisis. For EMEs, we notice that the effectiveness of macroprudential policies to lower the risk of a systemic banking crisis has significantly improved after the GFC; the coefficient estimate on the GFC-dummy variable is negative and significant at the 10 percent level in the EME sample in Table 7. This finding can be explained by a change in the type of macroprudential policies implemented in emerging market economies, by a change in the effectiveness of existing policies, or by a combination of both factors.

We confirm the statistically significant negative effect of real GDP growth on the conditional probability of a systemic banking crisis in all three cases. The economic significance of the positive effect of real economic activity on the crisis probability is quantitatively comparable to the results obtained in our baseline specification. For the linear crisis index, the results for the full sample of countries and for the sub-sample of advanced economies are also comparable to the baseline results. We note that, in the case of emerging market economies, the persistence of the linear crisis index increases and becomes statistically significant.

To summarize, accounting for potential changes after the GFC in the transmission of macroprudential policy to the conditional probability of a systemic banking crisis seems justified in the case of emerging market economies. This conclusion is also motivated by the better fit of this alternative model specification for the sub-sample of emerging market

Table 7: Macroprudential Policies and Banking Crises: the impact of the Global Financial Crisis

	All	AE	EME	All	AE	EME
MPI	-0.606^{***}	-0.411	-0.899^{***}	-0.535^{**}	-0.66^{*}	-0.506
	(0.175)	(0.292)	(0.286)	(0.252)	(0.383)	(0.377)
MPI $\times\ \mathbb{1}_{\{t-1\ \geq\ 2008\}}$				-0.061	0.239	-0.43^{*}
				(0.15)	(0.253)	(0.256)
GDP Growth	-0.192^{***}	-0.283^{***}	-0.18^{***}	-0.195^{***}	-0.264^{***}	-0.228^{***}
	(0.039)	(0.065)	(0.059)	(0.04)	(0.066)	(0.068)
Crisis Index	0.56^{***}	0.534^{***}	0.2	0.581^{***}	0.476^{***}	0.289^{*}
	(0.113)	(0.143)	(0.185)	(0.112)	(0.149)	(0.169)
FD	12.574^{***}	11.884^{***}	18.289^{**}	12.727^{***}	12.341^{***}	21.31^{***}
	(3.308)	(4.06)	(7.205)	(3.334)	(4.063)	(7.873)
Debt-to-GDP	-0.009	-0.026^{**}	0.013	-0.008	-0.03^{**}	0.005
	(0.007)	(0.012)	(0.015)	(0.007)	(0.012)	(0.016)
KA	-1.439	-2.058	-0.779	-1.572	-1.842	-1.867
	(1.109)	(1.595)	(1.725)	(1.135)	(1.64)	(1.954)
Trade-to-GDP	0.013	0.024	0.045	0.012	0.028	0.05
	(0.012)	(0.021)	(0.028)	(0.012)	(0.021)	(0.032)
AIC	532.396	321.042	170.112	531.382	325.101	165.986
BIC	727.937	429.619	238.075	731.368	437.556	237.347
Pseudo R^2	0.28	0.29	0.328	0.285	0.284	0.36
# Effect. Obs.	629	357	221	629	357	221
# Observations	2057	544	986	2057	544	986
Country FE	Y	Y	Y	Y	Y	Y
Year FE	N	N	N	N	N	N

Note: ML estimation is carried out using the iterative pseudo-demeaning algorithm of Stammann et al. (2016). The dependent variable is the systemic banking crisis binary. All regressors are considered as lagged values. We report bias-corrected estimates using the analytical bias correction of Hahn & Kuersteiner (2011). Standard errors clustered at the country level are reported between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We report the Akaike (AIC) and Bayesian (BIC) information criteria together with the McFadden pseudo R squared for each specification. The estimation period is 2001-2017 and we consider one-way country fixed effects. The total and effective number of observations considered in the estimation are also reported.

economies, both in terms of lower information criteria (AIC and BIC) and in terms of higher pseudo R^2 compared to the baseline specification.

We now turn our attention to the evaluation of the impact of the Global Financial Crisis on the trade-off between macroprudential regulation and real economic activity. The results of the estimation of the linear panel model specified in equation (4) are reported in the right panel of Table 8 below. Again, We report the results of our baseline specification (equation 2) in the left panel for ease of comparison.

Table 8: Macroprudential Policies and real GDP growth: the impact of the Global Financial Crisis

	All	AE	EME	All	AE	EME
MPI	-0.237^{***}	-0.037	-0.318^{***}	0.297^{**}	0.712^{***}	0.161
	(0.079)	(0.132)	(0.122)	(0.121)	(0.267)	(0.163)
MPI $\times \mathbb{1}_{\{t-1 \geq 2008\}}$				-0.434^{***}	-0.644^{***}	-0.419^{***}
				(0.075)	(0.199)	(0.092)
GDP Growth	0.306^{***}	0.368^{***}	0.372^{***}	0.282^{***}	0.352^{***}	0.336^{***}
	(0.022)	(0.044)	(0.031)	(0.022)	(0.044)	(0.032)
Crisis Index	-0.204^{**}	-0.202^{*}	-0.147	-0.179^{**}	-0.152	-0.203^{*}
	(0.092)	(0.117)	(0.150)	(0.089)	(0.123)	(0.119)
FD	-6.320^{***}	-6.863^{**}	-5.510^{**}	-4.541^{**}	-6.569^{**}	-2.885
	(1.792)	(2.977)	(2.446)	(1.802)	(2.954)	(2.476)
Debt-to-GDP	0.011^{***}	0.023^{***}	0.028^{***}	0.010^{***}	0.034^{***}	0.023^{***}
	(0.004)	(0.009)	(0.008)	(0.004)	(0.009)	(0.008)
KA	-0.541	0.969	-0.766	-0.846	0.688	-1.135
	(0.627)	(1.379)	(0.778)	(0.624)	(1.367)	(0.773)
Trade-to-GDP	0.002	-0.031^{***}	0.021^{**}	0.002	-0.034^{***}	0.021^{**}
	(0.005)	(0.011)	(0.009)	(0.005)	(0.011)	(0.009)
Country FE	Y	Y	Y	Y	Y	Y
Year FE	N	N	N	N	N	N
Observations	2,057	544	986	2,057	544	986
R^2	0.361	0.299	0.347	0.372	0.315	0.362
Adjusted R^2	0.319	0.247	0.301	0.331	0.262	0.317

Note: Estimation is carried out by OLS with one-way country fixed effect dummies. The estimation period is 2001-2017. The dependent variable is the year-on-year real GDP growth rate. All regressors are considered as lagged values. We report estimates with Standard errors clustered at the country level between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We report both the R-squared and adjusted R-squared for each specification.

A first interesting result is that the effect of macroprudential regulation on real economic activity appears to have changed substantially after the GFC. Indeed, the MPI coefficient has a statistically significant positive sign when considering all the countries pooled together and the sub-sample of advanced economies. If we consider, on the other hand, the interaction variable between the MPI variable and a post-crisis indicator variable, the coefficient estimates are negative and statistically significant in all the cases considered. We thus confirm our baseline result that macroprudential policies depress real economic activity for the post-GFC period when considering all the countries pooled together and emerging market economies.

Our alternative model specification does not alter the persistence or statistical significance of real GDP growth. Interestingly, the pervasiveness of the adverse impact of systemic banking crisis on real economic activity, measured by the coefficient estimate for the linear crisis index, becomes statistically significant in the case of EMEs while the opposite is true for AEs. The statistically significant negative impact of financial development on real economic activity is confirmed in all cases but emerging market economies. The statistical significance of Debt-to-GDP (positive sign) and Trade-to-GDP (negative sign for AEs and positive sign for EMEs) is also confirmed.

To conclude this section, our alternative specification seems also relevant in the case of the modeling of real GDP growth for two reasons. First, the sensitivity of real economic activity to macroprudential regulation seems to have experienced an important change in the period following the Global Financial Crisis. This is documented by the opposite signs of the coefficient for MPI and for the interaction variable between the MPI variable and a post-crisis indicator variable. Second, this alternative specification seems supported by the data, as indicated by the higher values of the adjusted R^2 in all three cases.

6. Conclusion

This paper uses a novel empirical setup to assess the effectiveness of macroprudential policies in achieving their ultimate goal of reducing the incidence of financial crises. Our empirical specification allows us to quantify the effect of these policies on the probability of financial crises and on economic growth. Our main objective is to understand the channels through which macroprudential policies can have an influence on financial stability. In particular, we identify two possible channels. First, macroprudential policies can have a direct effect on the probability of banking crises if they happen to smooth financial cycles. Second, macroprudential policies can exert an indirect effect on the probability of banking crises by weakening economic growth. To this end, we use an empirical method that allows us to disentangle the direct and indirect effects of macroprudential policies as well as their net effect on the probability of systemic banking crises.

Our results indicate that macroprudential policies have a statistically and economically direct negative effect on the probability of banking crises. Moreover, we find that macroprudential policies depress economic growth. However, when we account for these two effects together, we find that macroprudential policies have a positive net effect on financial stability; even though, MPPs lower economic growth, they still lead to a lower incidence of systemic banking crises. Furthermore, our results suggest that the mitigating effect of MPPs on the likelihood of banking crises is more pronounced in emerging market economies relative to advanced economies. Our estimations also suggest that borrower-based macroprudential tools seem to be more effective in emerging market economies while financial-based macroprudential tools are more useful in taming the incidence of banking crises in advanced economies.

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Appendix A. Additional information on data

Table A.9: List of countries used in the empirical analysis and group classification

Advanced Economies (AE)	Emerging Market Economies (EME)	Low-Income Developing Countries (LIDC)
Australia	Albania, Lebanon	Bangladesh
Austria	Algeria, Malaysia	Bhutan
Belgium	Angola, Mauritius	Burkina Faso
Canada	Argentina, Mexico	Cabo Verde
Cyprus	Armenia, Mongolia	Cambodia
Czech Republic	Azerbaijan, Morocco	Côte d'Ivoire
Denmark	Belarus, Namibia	Ghana
Estonia	Belize, Nigeria	Guinea-Bissau
Finland	Bolivia, North Macedonia	Guyana
France	Bosnia and Herzegovina	Haiti
Germany	Botswana, Pakistan	Honduras
Greece	Brazil, Panama	Kenya
Iceland	Bulgaria, Paraguay	Kyrgyz Republic
Ireland	Chile, Peru	Lao P.D.R.
Israel	China, Philippines	Madagascar
Italy	Colombia, Poland	Mali
Japan	Costa Rica, Russia	Moldova
Korea	Croatia, South Africa	Mozambique
Latvia	Dominican Republic	Myanmar
Lithuania	Ecuador, Sri Lanka	Nepal
Netherlands	Egypt, St. Kitts & Nevis	Nicaragua
New Zealand	El Salvador, Thailand	Niger
Norway	Georgia, Tunisia	Rwanda
Portugal	Guatemala, Turkey	Senegal
Singapore	Hungary, Ukraine	Sierra Leone
Slovak Republic	India, Uruguay	Tajikistan
Slovenia	Indonesia, Vietnam	Tanzania
Spain	Islamic Republic of Iran	The Gambia
Sweden	Jamaica	Togo
Switzerland	Jordan	Uganda
United States	Kazakhstan	Zambia
United Kingdom	Kuwait	

Appendix B. GMM estimation for macroprudential policies and real GDP growth

Table B.10 below provides the estimates for the model specified in Equation (2) using the GMM estimator introduced in Arellano and Bond (1991). The OLS estimates from Table 4 are also reproduced for ease of comparison.

The reported p-values for a Sargan test on the joint validity of the instruments do not provide evidence supporting a rejection of the null hypothesis on the adequacy of the selected instruments.[15] We confirm our main result that tighter macroprudential regulation depresses economic growth. The estimated impact of the MPI on real economic activity is larger with the GMM estimator than with the OLS approach. For instance, a one unit increase in the MPI index lowers economic growth by 79.8 basis points in the full sample when considering the GMM results and 23.7 basis points for the OLS results. In addition, the estimated coefficient for the subsample of AEs is now statistically significant and its magnitude is comparable to the one for EMEs. We also confirm our results for the negative and significant impact of the banking crisis index in the full sample and in the subsample of AEs. We observe that the magnitude of the estimated coefficients is approximately doubled when comparing the GMM estimates to the OLS ones. Concerning the macroprudential controls, we confirm that countries with higher public debt relative to GDP enjoy higher economic growth and trade openness depresses economic activity in both AEs and EMEs.

[15] Note that it only holds at a 1% significance level in the subsample of EMEs.

Table B.10: Macroprudential Policies and Real GDP growth (GMM results)

	All		AE		EM	
	OLS	GMM	OLS	GMM	OLS	GMM
MPI	-0.237^{***}	-0.798^{***}	-0.037	-0.614^{**}	-0.318^{***}	-0.550^{*}
	(0.079)	(0.199)	(0.132)	(0.297)	(0.122)	(0.286)
GDP Growth	0.306^{***}	0.294^{***}	0.368^{***}	0.641^{***}	0.372^{***}	0.414^{***}
	(0.022)	(0.051)	(0.044)	(0.120)	(0.031)	(0.080)
Crisis Index	-0.204^{**}	-0.464^{**}	-0.202^{*}	-0.567^{**}	-0.147	-0.199
	(0.092)	(0.236)	(0.117)	(0.281)	(0.150)	(0.490)
FD	-6.320^{***}	-1.208	-6.863^{**}	6.093	-5.510^{**}	6.635
	(1.792)	(5.062)	(2.977)	(5.623)	(2.446)	(6.267)
Debt-to-GDP	0.011^{***}	0.06^{***}	0.023^{***}	0.188^{***}	0.028^{***}	0.141^{***}
	(0.004)	(0.016)	(0.009)	(0.037)	(0.008)	(0.023)
KA	-0.541	-5.044^{*}	0.969	5.208	-0.766	-2.763
	(0.627)	(2.956)	(1.379)	(3.528)	(0.778)	(1.899)
Trade-to-GDP	0.002	-0.031^{**}	-0.031^{***}	-0.181^{***}	0.021^{**}	-0.053^{**}
	(0.005)	(0.014)	(0.011)	(0.049)	(0.009)	(0.022)
Country FE	Y	Y	Y	Y	Y	Y
Year FE	N	N	N	N	N	N
Observations	2,057	1815	544	480	986	870
Sargan Test	–	0.109	–	0.275	–	0.021

Note: The model estimated is specified in Equation (2). The estimation period is 2001-2017. The dependent variable is the year-on-year real GDP growth rate. All regressors are considered as lagged values. The OLS columns reproduce the estimates reported in Table 4. The GMM columns report the estimates from the Arellano and Bond (1991) GMM estimator. In the sample with all countries, we treat all the right-hand side variables as endogenous in the estimation. For the subsamples of AEs and EMEs, we treat GDP growth and the crisis index as endogenous and the remaining variables as additional instruments to avoid the issues arising from the use of numerous instruments in dynamic panel GMM estimation with a small cross-sectional dimension. We report estimates with Standard errors clustered at the country level between brackets. ***, **, and * indicate significance at the 1-, 5-, and 10-percent significance levels, respectively. We also report the p-values for the Sargan test of joint validity of the instruments.

Appendix C. Generalized Impulse Response Function (GIRF) of a tightening shock to macroprudential supervision

We detail the procedure used to compute the response of the crisis probability and of real GDP growth to a (exogenous) shock to the macroprudential index. In the spirit of the Generalized Impulse Response Function (GIRF) approach, we do the following steps for $h = 0, \ldots, H$:

- For $h = 0$: we initialize all the explanatory variables to their unconditional means.[16] The country fixed effect parameters from equations (1) and (2) are set to the mean of the estimated parameters in each case.
- We compute 2 paths of the bivariate system consisting of the conditional probability of crisis and real GDP growth: the baseline case where variables are initialized as explained above and the second where we add a one-period shock, $\delta > 0$, to the initial value of the macroprudential index (MPI).
- For $h > 0$: the exogenous variables are set to their initial (unconditional mean) value. The endogenous variables (real GDP growth and the latent crisis index) are updated recursively using their lagged values according to our two-equation system (see equations (1) and (2)).
- The GIRF for the crisis probability and real GDP growth are obtained as the difference between the path with the one-period macroprudential tightening (i.e., positive shock) and the baseline path.

[16] We first compute country-specific averages and then take a cross-sectional average.

Appendix D. Moving-block bootstrap procedure for GIRFs

We repeat, for clarity, the equation for the dynamic panel logit with fixed effect:

$$\begin{aligned} y_{1,it} &= \mathbb{1}_{\{y^*_{1,it} \geq 0\}} \\ y^*_{1,it} &= \alpha_{1,i} + \mathbf{x}^\top_{it-1}\ \boldsymbol{\beta}_1 + \mathbf{y}^\top_{it-1}\ \boldsymbol{\rho}_1 + u_{1,it} \end{aligned} \tag{D.1}$$

Where we made a distinction between exogenous ($\mathbf{x}_{it}$) and endogenous variables ($\mathbf{y}_{it}$) of our dynamic system. The second equation for real GDP growth ($y_{2,it}$) takes the form of a dynamic linear panel with fixed effects:

$$y_{2,it} = \alpha_{2,i} + \mathbf{x}^\top_{it-1}\ \boldsymbol{\beta}_2 + \mathbf{y}^\top_{it-1}\ \boldsymbol{\rho}_2 + u_{2,it} \tag{D.2}$$

Given parameter estimates $\hat{\boldsymbol{\theta}}_1 \equiv \left[\hat{\boldsymbol{\alpha}}^\top_1,\ \ \hat{\boldsymbol{\beta}}^\top_1,\ \ \hat{\boldsymbol{\rho}}^\top_1\right]^\top$ for the dynamic panel logit model in equation (D.1) and $\hat{\boldsymbol{\theta}}_2 \equiv \left[\hat{\boldsymbol{\alpha}}^\top_2,\ \ \hat{\boldsymbol{\beta}}^\top_2,\ \ \hat{\boldsymbol{\rho}}^\top_2\right]^\top$ for the dynamic linear panel model in equation (D.2), we define a new matrix $\boldsymbol{Z}_t \equiv \left[\hat{\boldsymbol{u}}_{1t},\ \ \hat{\boldsymbol{u}}_{2t},\ \ \boldsymbol{X}_{t-1}\right]$ which maintain the cross-sectional dependence between the I countries for the fitted residuals ($\hat{u}_{1,it}$ and $\hat{u}_{2,it}$) and for the K exogenous variables indexed by $\mathrm{x}_{k,it-1}$:

$$\boldsymbol{Z}_t \equiv \begin{bmatrix} \hat{u}_{1,1t} & \hat{u}_{2,1t} & \mathrm{x}_{1,1t-1} & \mathrm{x}_{2,1t-1} & \dots & \mathrm{x}_{K,1t-1} \\ \hat{u}_{1,2t} & \hat{u}_{2,2t} & \mathrm{x}_{1,2t-1} & \mathrm{x}_{2,2t-1} & \dots & \mathrm{x}_{K,2t-1} \\ \vdots & \vdots & \vdots & \vdots & \vdots & \vdots \\ \hat{u}_{1,It} & \hat{u}_{2,It} & \mathrm{x}_{1,It-1} & \mathrm{x}_{2,It-1} & \dots & \mathrm{x}_{K,It-1} \end{bmatrix}$$

Note that in the case of countries which did not experience any systemic banking crisis over the period under consideration, and are therefore not taken into account in the estimation of model (D.1), we replace the fitted residuals of the model by a missing value.

We can now introduce the approach for the moving-block bootstrap (MBB). Given a block length b and a time-series dimension T for the panel, we can generate $T - b + 1$ overlapping blocks as follows:

$$\underbrace{\boldsymbol{Z}_1, \ \dots \ , \boldsymbol{Z}_b}_{\text{1 block of length } b}; \ \boldsymbol{Z}_2, \ \dots \ , \boldsymbol{Z}_{b+1}; \ \dots \ ; \ \boldsymbol{Z}_{T-b+1}, \ \dots \ , \boldsymbol{Z}_T$$

The algorithm to compute the bootstrapped confidence intervals for the GIRFs of our dynamic system consists in applying the following steps:

1) Draw (with replacement) $(T_{\text{burn}} + T)/b$ blocks from the $T - b + 1$ overlapping blocks where T_{burn} is a burn in period. Call the resulting sample replicate $\boldsymbol{Z}^{(r)}$

2) Compute recursively the model implied endogenous variables $y_{1,it}^{(r)}$ and $y_{2,it}^{(r)}$ using equations (D.1)-(D.2) with the respective estimated parameters $\hat{\boldsymbol{\theta}}_1$ and $\hat{\boldsymbol{\theta}}_2$. Keep only the last T observations to avoid that the bootstrapped samples have dependence on the initial conditions.

3) Re-estimate models (D.1)-(D.2) and store $\hat{\boldsymbol{\theta}}_1^{(r)}$ and $\hat{\boldsymbol{\theta}}_2^{(r)}$.

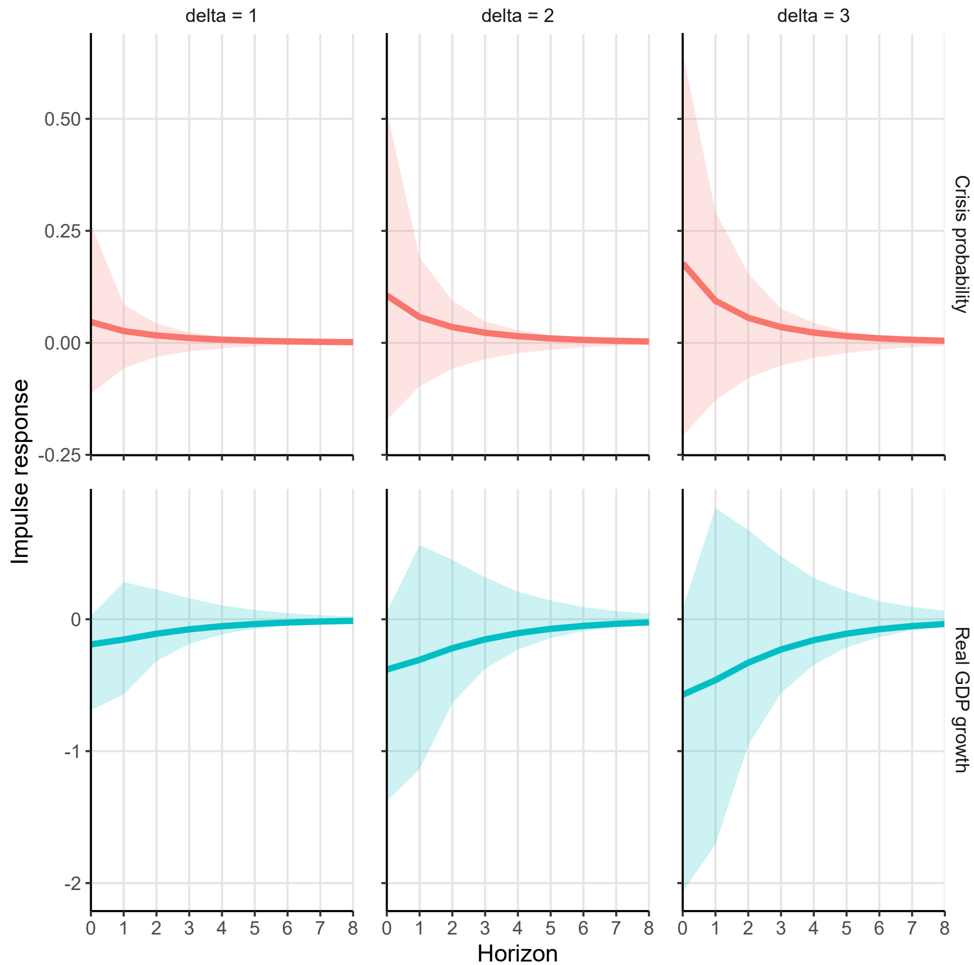
4) Compute and store the implied GIRFs using the approach outlined in Appendix C, $\mathcal{I}_{\text{Y}}(H, \delta, \Omega_{t-1})^{(r)}$.

5) Repeat points 1-4 for $r = 1, \ \dots \ , R$ where R is the total number of bootstrap replications.

6) For $h = 0, 1, \ \dots \ , H$, construct a 68% confidence interval around the GIRF for horizon h by taking the $16-$ and $84-$percentile of the distribution of the R replicas of $\mathcal{I}_{\text{Y}}(h, \delta, \Omega_{t-1})^{(r)}$.

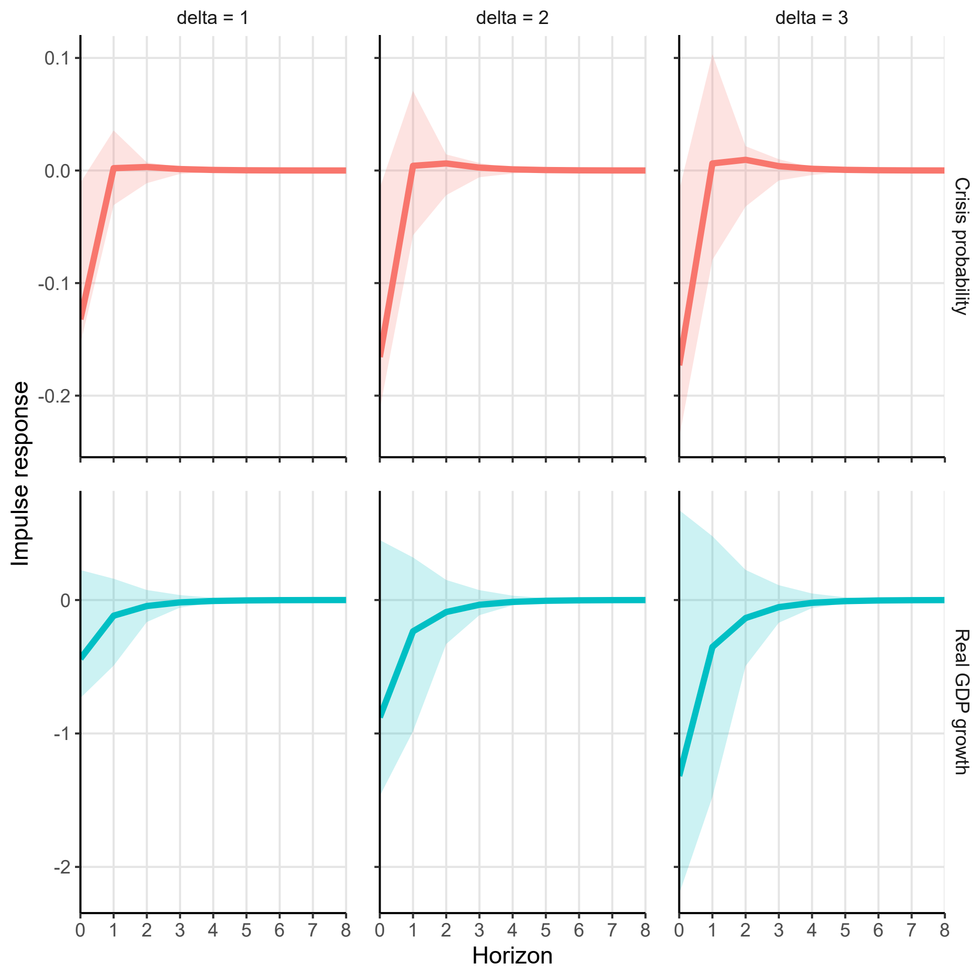
Appendix E. Targeting borrowers vs. financial institutions: GIRFs for MPI-Borrower

Figure E.5: GIRF analysis of a borrower macroprudential tightening: advanced economies



Note: Comparison of the impulse responses of the probability of a systemic banking crisis (top-panel in red) and of real GDP growth (bottom-panel in green) to different levels of borrower macroprudential tightenings (positive MPI-Bor shocks denoted by delta). The horizon of the responses displayed ranges up to eight years after the policy tightening. The Generalized Impulse Response Functions are computed following the procedure outlined in Appendix C. The shaded areas around the response functions correspond to the 68% confidence intervals from the Moving-Block Bootstrap procedure detailed in Appendix D with 5000 bootstrap replications and a block size of 5 years.

Figure E.6: GIRF analysis of a borrower macroprudential tightening: emerging market economies



Note: Comparison of the impulse responses of the probability of a systemic banking crisis (top-panel in red) and of real GDP growth (bottom-panel in green) to different levels of borrower macroprudential tightenings (positive MPI-Bor shocks denoted by delta). The horizon of the responses displayed ranges up to eight years after the policy tightening. The Generalized Impulse Response Functions are computed following the procedure outlined in Appendix C. The shaded areas around the response functions correspond to the 68% confidence intervals from the Moving-Block Bootstrap procedure detailed in Appendix D with 5000 bootstrap replications and a block size of 5 years.