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Assessing the fiscal implications of banking crises

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We propose a method for computing the distribution of the potential fiscal cost of a banking crisis – a key input in assessing the adequacy of a country's fiscal buffers. First, we use a cross-section of banking crises to identify the risk factors that predict the post-crisis increase in public sector debt – a measure of the overall fiscal cost of a crisis. Next, we use these risk factors to compute country-specific distributions of that cost in the event of a crisis. We find that the level and growth of credit to the private non-financial sector, foreign exchange reserves and the ratio of bank capital to assets are relevant predictors. As an illustration, we apply the method to the conditions prevailing in 2018 and find that the potential fiscal costs could be sizeable: with a probability of 95%, public debt could approach 40% of GDP on average across countries. Our illustrative estimates are probably upper bounds: while they indicate that higher bank capital can substantially reduce fiscal costs, they exclude the broader benefits of the wide-ranging reforms after the Great Financial Crisis.

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Keywords: banking crises; public debt; fiscal space; fiscal buffers; macro-financial stability framework.

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Introduction

One of the most significant and enduring legacies of a banking crisis is a large increase in public debt (eg Reinhart and Rogoff (2009, 2013), Laeven and Valencia (2013, 2018), Bova et al (2016) and Furceri and Zdzienicka (2012)). This increase stems from two sources. First, fiscal resources are spent to recapitalise and restructure failed financial institutions (direct costs). Second, debt surges as deficits soar due to the drop in economic activity and, where feasible, the countercyclical policy response to stabilise the economy (indirect costs). On average, the indirect costs have been larger than the direct costs (Reinhart and Rogoff (2009, 2011a), Laeven and Valencia (2013, 2018)).[1]

Large crisis-induced surges in public debt may, in turn, have severe consequences.

Ex post, the sovereign could lose normal market access or even face insolvency. Bordo and Meissner (2016)) document how banking crises followed or accompanied by fiscal crises are rarer but more costly in terms of output losses.[2] Even if a fiscal crisis is avoided, the large increase in public debt may cripple the government's ability to run a countercyclical fiscal policy, slowing the recovery. Moreover, higher public debt may narrow the room for future countercyclical policy and weaken trend growth.[3]

Ex ante, financial market concerns with public debt can themselves contribute to a crisis. The expectation that fiscal resources may not suffice to backstop the financial system and the economy can generate financial instability. The cost and availability of funding may tighten sharply both for the sovereign and for the entire economy, sapping economic activity and destabilising financial institutions. Adverse loops may arise, in which fiscal and financial risks feed on each other. Such loops were clearly at

1 For a discussion of budgetary risks, see also Eschenbach and Schuknecht (2004).

2 A government that loses market access may resort to money printing to service its domestic currency debt, but the cost of the resulting likely high inflation may not necessarily be smaller than debt restructuring. Indeed, Jeanneret and Souissi (2016) find that high inflation *increases* the probability of default on local currency debt. Reinhart and Rogoff (2011b) show that default on domestic currency debt has historically been more frequent than commonly thought. Over the past decades, with financial liberalisation and the deepening of domestic capital markets in many emerging market economies, the incidence of local currency defaults has increased. Out of 63 sovereign defaults since 1975, 26 were foreign currency debt defaults, 20 were local currency defaults only and 14 involved both domestic and foreign currency bonds (three are still unresolved) (Moody's (2019)). As a result, the ratings gap between domestic and foreign currency debt has diminished (Amstad et al (2018)). With rare exceptions, Moody's currently assigns the same rating to both types of debt (Moody's (2019)).

3 Although private credit booms rather than public debt increases tend to precede and cause financial crises (eg Boissay et al (2016), Drehmann et al (2012), Gorton and Ordoñez (2019), Jordà et al (2013), Schularick and Taylor (2012)), higher public debt is likely to make financial crises more costly in terms of output. Jordà et al (2016) find evidence for advanced economies that higher public debt at the onset of a financial crisis leads to deeper recessions and slower recoveries (Jordà et al (2016)). Romer and Romer (2018) find that output costs are much bigger when a country has little or no monetary and fiscal space before a financial crisis. Aizenman et al (2019) show that, on average, countries that have a more limited fiscal capacity follow a more procyclical fiscal policy. High government debt may also weigh on trend output growth (eg Cecchetti et al (2011), Checherita-Westphal and Rother (2012), Baum et al (2013), Woo and Kumar (2015), Chudik et al (2016)). See also the review in Reinhart et al (2012). The causal link is likely to run in both directions (eg Panizza and Presbitero (2014)).

work in a number of euro area countries in 2010–12,[4] but they are not unique to countries sharing a common currency. For example, countries heavily reliant on foreign funding might experience large currency depreciations and tighter financial conditions when investors question fiscal sustainability.[5] This may be true even for countries without large amounts of foreign currency debt but with a large presence of foreign investors in local currency debt (eg Carstens and Shin (2019)). In these countries, a weak fiscal position could limit the scope of monetary policymakers to respond countercyclically.

These considerations put a premium on assessing the adequacy of a country's fiscal buffers against banking crisis risk.[6] Unfortunately, this is not a straightforward task. The fiscal cost of a crisis varies greatly across episodes and countries, making historical averages or medians imperfect measures of the potential cost. Understanding what factors explain this variation should inform estimates of financial contingent liabilities and of the corresponding buffers. Yet, while a large literature has assessed the probability of financial crises, relatively few studies have analysed their fiscal cost (see below). For a decision-maker, estimating the losses from a rare event such as a crisis should be at least as important as estimating its probability.[7]

In this paper, we take a step towards filling this gap. We investigate what factors can explain the fiscal cost of a financial crisis and develop a method for estimating the ex ante distribution of this cost as a function of prevailing economic conditions. Purely as an illustration, we then apply it to conditions in 2018.

Specifically, we proceed in three steps. First, we follow the literature on financial crises (eg Laeven and Valencia (2013, 2018), Bova et al (2016)) and approximate the overall ex post fiscal cost (ie the cost given a crisis) with the increase in (gross) public debt over a five-year window.[8] Such a measure includes both the direct costs and the,

4 Studies find that during the euro area debt crisis, sovereign risk widened corporate spreads, depressed bank valuations and curtailed lending (eg Altavilla et al (2017), Bottero et al (2016), Correa et al (2014), Demirgüç-Kunt and Huizinga (2013), Popov and van Horen (2015), Bocola (2016)), thereby weakening economic activity. Similarly, higher public debt and fiscal deficits raised the correlation between sovereign and bank CDS spreads, indicating that fiscal and financial risks reinforced each other (eg Acharya et al (2014), de Bruyckere et al (2013), Aizenman et al (2013), Kallestrup et al (2016)).

5 This may occur even if *current* fiscal deficits and debt are not considered excessive. One of the factors behind the financial crises experienced in a number of EMEs in the 1980s and 1990s, including the 1997 Asian crisis, was the expectation of a future bailout (and the ensuring unsustainable future fiscal policy); see eg Velasco (1987), Calvo and Mendoza (1996) and Corsetti et al (1999)).

6 They also put a premium on preserving fiscal buffers in good times and, where buffers are judged insufficient, on rebuilding them For example, the IMF (2018) stresses the importance of taking advantage of cyclical upswings to rebuild fiscal buffers and be in a better position to weather future downturns and unexpected rare events, including financial crises. See also Obstfeld (2013).

7 In addition, existing estimates of fiscal space are highly uncertain and sensitive to the method used. Also, such estimates do not generally incorporate either implicit liabilities from future pension and health spending or financial crisis-related contingent liabilities. See Box V.B in BIS (2016) and Corsetti (2018) for a discussion. Ganiko et al (2016) account explicitly for uncertainty, reaching the conclusion that several EMEs have little fiscal space.

8 Our focus on the cost of a crisis is akin in credit risk management to the analysis of loss-given-default as opposed to the probability of default. For a definition of these concepts, see eg McNeil et al (2005).

on average larger, indirect costs.[9] Second, we identify the relevant risk factors by linking the ex post fiscal cost to the behaviour of several variables in a sample of financial stress episodes from 1976 to 2012.[10] Third, we compute the mean and various percentiles of the conditional distribution of the fiscal cost given a crisis based on quantile regressions that include the risk factors identified in the previous step. To approximate this distribution we use an interpolation method akin to that used in Adrian et al (2019).

Our main findings can be summarised as follows.

First, the level and growth rate of domestic credit to the private non-financial sector, are economically significant predictors of the fiscal cost of crises, as are those of foreign exchange reserves and bank capital.

Second, the fiscal buffer – the amount of fiscal space needed to cover an unexpected increase in public debt after a crisis while avoiding a sovereign default or loss of normal market access – can be sizeable and may vary substantially across countries. As an illustration, conditional on a crisis occurring, public debt could increase, on average across advanced and emerging market economies, to approach 40% of GDP with a 95% probability.[11,12]

Finally, our findings regarding aggregate bank capital underline the importance of prudential regulation to reduce the fiscal costs. Our illustrative counterfactual analysis suggests that, if countries had the same bank capital in 2008 as in 2017, the mean fiscal cost following the GFC would have been about 7 percentage points of GDP lower, on average, all else equal. In the countries that incurred the worst costs, such as Ireland, the gain would have been much larger. Moreover, our estimates cannot take into account the benefits of the much more wide-ranging post-crisis financial regulatory reforms (Borio et al (2020)).

Some words of caution in interpreting our results are warranted. First, any estimate of a fiscal cost is inevitably subject to uncertainty due to the method used, data availability and the impossibility of incorporating all relevant information. For this reason, our methodology is intended to be just one input in what is a much more complex decision. In particular, our methodology does not take into account the

9 Because we measure the change over a window of a few years after the start of a crisis, it is less likely that this measure understates the true fiscal cost. First, estimates of the direct bail out fiscal costs may not take into account recovery rates. But after a few years any recovery would be recorded as a reduction in the debt stock, other things equal. Second, any initial discretionary fiscal stimulus that may be needed at the height of the crisis may be subsequently reversed, at least partly.

10 In selecting these variables, we take inspiration from the literature on financial crises (eg Kaminsky and Reinhart (1999), Borio and Lowe (2002), Borio and Drehmann (2009, 2011), Schularick and Taylor (2012), Gourinchas and Obstfeld (2012)). That said, we have no strong a priori reasons to believe that the set of factors that influence the probability of a crisis should also affect its cost and vice versa.

11 These estimates are well within the historical record. Given our measure of fiscal costs, public sector debt increases on average by 22% of GDP within a five-year window. Laeven and Valencia (2018) report a mean increase of 21% for high-income countries and 16% for low- and middle-income countries within a four-year window. In extreme cases, they report increases of public debt of more than 70% of GDP for high-income countries and more than 100% of GDP for low- and middle-income countries. We find similar values within our preferred time window.

12 Conventional estimates of fiscal space (eg Fournier and Fall (2015), Botev et al (2016), OECD (2016)) suggest that most countries have enough fiscal space to absorb a financial crisis. However, as noted in footnote 7, such estimates are highly uncertain.

probability of a crisis; only the fiscal cost given a crisis – akin to what in risk management is the loss-given-default. Second, our illustrative estimates might not capture the full benefits of the financial reforms following the Great Financial Crisis (GFC). These have not just raised bank capital, but also improved its quality and robustness, introduced liquidity standards, implemented macroprudential frameworks, and put in place specific arrangements to ensure the orderly resolution of systemically important banks (BIS (2018), Borio et al (2020)). Finally, if the potential fiscal cost of a crisis is large, the best solution does not necessarily or exclusively lie in fiscal measures. Depending on a country's specific circumstances, tighter prudential regulation could be important in normal times to increase the resilience of the financial sector.

Our paper contributes to the literature on the fiscal consequences of financial crises (eg Reinhart and Rogoff (2009, 2013), Laeven and Valencia (2013, 2018), Bova et al (2016), Furceri and Zdzienicka (2012)). Using the post-crisis change in public debt as a measure, these studies report simple statistics regarding the distribution of fiscal costs, focusing, for instance, on country subgroups or subperiods. However, they do not systematically investigate the factors behind the variation of fiscal costs across episodes.[13] To the best of our knowledge, Amaglobeli et al (2017) and ours are the only studies that have modelled this variation. Our study differs from Amaglobeli et al (2017) in that we use a larger set of countries and crises as well as a different econometric specification. Another important difference is that we also compute the distribution of potential fiscal cost using a measure akin to the value-at-risk. The measure is routinely used in risk management and more recently it has also been employed in macroeconomic analysis to measure GDP-at-risk (Adrian et al (2019)).

Our analysis complements previous findings in the literature on the macroeconomic effects of financial crises. In particular, in a sample of advanced and emerging market economies, Cecchetti et al (2011) find that a lower level and growth rate of credit to the private non-financial sector, larger aggregate bank capital and larger foreign exchange reserves go hand in hand with better macroeconomic performance in the aftermath of the GFC. Berkmen et al (2012) obtain similar findings. Bank capital also appears to matter based on long-run data. In a study that covers advanced economies over the period 1870–2013, Jordà et al (2017) find that, while pre-crisis bank capital does not help predict a crisis, it reduces its output cost substantially. Specifically, depending on whether capital is above or below the historical average, the difference in real GDP per-capita is 5 percentage points five years after the start of the crisis. These results are consistent with our finding that bank capital significantly reduces the fiscal cost of a banking crisis.

The rest of the paper is organised as follows. Section 1 introduces the data and the model specification. Section 2 describes the estimation results and identifies the risk factors. Section 3 describes how to compute the predictive distribution of fiscal costs given a crisis and, as an illustration, presents the computation for a selected group of advanced and emerging market economies. Finally, the conclusion highlights some policy implications of our findings.

[13] In addition to financial crises, countries may require specific fiscal space to deal with other uncertain contingencies such as natural catastrophes, subnational bankruptcies or even wars. Bova et al (2016) document that the associated fiscal costs can be very large, approaching that of financial events.

1. Data and model specification

To construct a cross-section of financial crisis episodes, we proceed as follows. We take the dating of banking crises in the recently compiled data set of Bova et al (2016). This large data set includes the systemic banking crises identified by Laeven and Valencia (2013, 2018), the non-systemic crises from Honohan and Klingebiel (2000) and other episodes identified by examining data on stock-flow adjustments from Eurostat (2012).[14] We then compute the five-year cumulative increase in gross government debt (as a percentage of GDP) from the start of a crisis (ie the change in debt between t–1 and t+4, where t indicates the crisis year).

The main issue with this measure is that it is negative in a number of episodes. Closer inspection reveals that most of these are cases in which the government defaulted: that is, the decline in debt after financial stress reflects debt restructuring or write-offs. For these episodes, we use an alternative proxy: the cumulative increase in the primary fiscal deficit over the five-year window. This can be regarded as a lower bound for the true fiscal cost. The remaining episodes in which debt declines are: Switzerland 2008; Norway 2009, Estonia 1998, Kuwait 2008, Macedonia 2005 and Sweden 2008. As these observations do not correspond to an increase in debt, or we lack the data on the primary balance for those episodes, we drop them from our sample.[15] Importantly, the costless nature of some of these crises reflects to a large extent the fact that banks incurred losses on foreign exposures while the domestic economy held up relatively well. Using a procedure to detect sample selection bias (Heckman (1979)), we conclude that dropping them does not affect our results (Section 2 and Annex B).

We complete our data set by including the predetermined values (at t–1 or earlier) of a set of potential explanatory variables (to be described later). We end up with a sample of 67 episodes covering the period 1976–2012 (the year of the last crisis included). The sample covers 25 advanced economies and 42 emerging market and less developed economies.

Our measure of overall fiscal cost appears to be well approximated by a log-normal distribution (Figure 1). Moreover, statistical tests (not reported) fail to reject the null that the log of the fiscal cost is normally distributed. This suggests using a log-level specification:

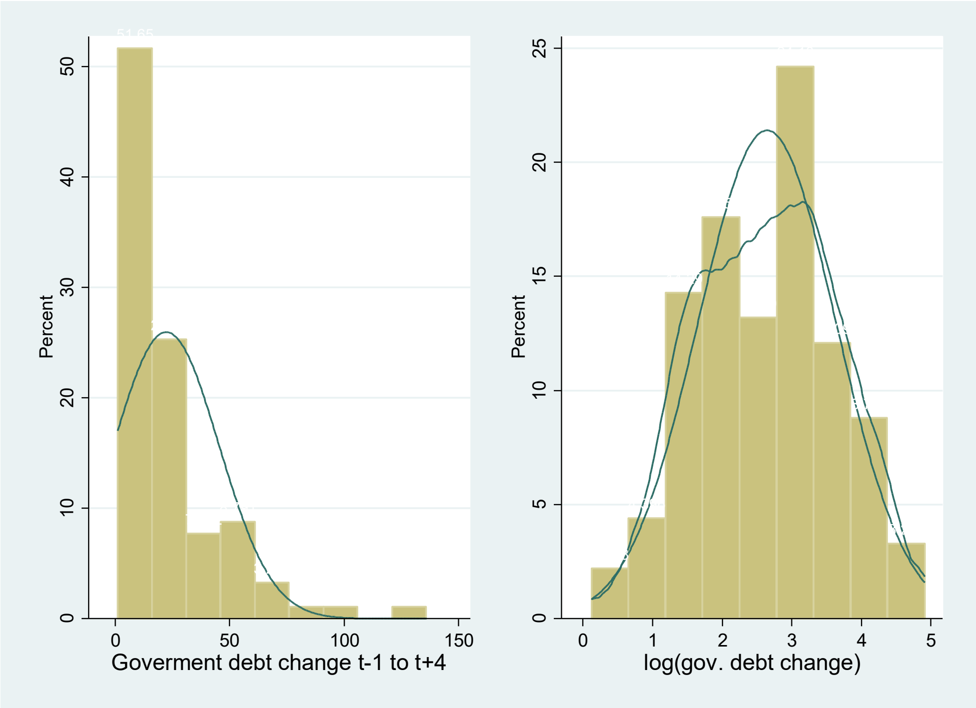
$$y_i \equiv \ln(c_i) = a + bX_i + \epsilon_i \qquad (1)$$

[14] Laeven and Valencia (2013) define systemic banking crises as periods with: (*i*) significant signs of financial distress in the banking sector (significant bank runs, losses in the banking system and/or bank liquidations); and (*ii*) significant banking policy intervention in response to losses in the banking system. For more details, see Laeven and Valencia (2013). Honohan and Klingebiel (2000) date episodes drawing on Caprio and Klingebiel (1997), Caprio and Klingebiel (1999), Lindgren et al (1996) and from consultations with countries' experts.

[15] If we choose a shorter time window of three years, we obtain a sample with fewer non-defaulters that experience a fall in debt. But we would underestimate the overall fiscal cost for the majority of countries.

where $c_i > 0$ is the fiscal cost, X_i is a set of predetermined explanatory variables (or risk factors) and ϵ_i is a set of iid errors. To estimate (1) we can use OLS. To obtain consistent estimates of b, we do not need to assume that the errors are normally distributed; thus, we adopt standard procedures to correct the standard errors for possible heteroscedasticity. In any case, in the specifications in Section 2 we could not reject the null hypothesis that the residuals are normally distributed.[16]

The distribution of the overall fiscal cost of banking crises — Figure 1



Note: overall fiscal cost is proxied by the increase in gross public debt in the period between one year before and four years after the onset of a banking crisis. The graph shows a log-normal distribution in the left-hand panel and a normal distribution and a kernel approximation in the right-hand panel.

Source: Bova et al (2016); authors' calculations.

2. Risk factors

We consider a broad set of factors that may plausibly influence the overall fiscal cost of a banking crisis. These include: the leverage of the private non-financial sector (measured by credit to the domestic non-financial sector as a ratio to GDP); the size and leverage of the financial sector (the ratio of bank assets to GDP, the ratio of loans to deposits, and the ratio of bank capital to assets, risk-weighted or not); the

[16] The analysis of residuals showed two big outliers, Brazil 2002 and Jordan 2004, which we dropped from the sample. These are excluded from Figure 1.

economy's pre-crisis cyclical position (a one-sided HP-filtered output gap,[17] pre-crisis average growth and the pre-crisis commodity price increase); the fiscal position (government debt or primary fiscal balance as a share of GDP); the country's external position (eg the current account balance, net foreign asset position, real exchange rate measures and foreign exchange reserves as a ratio to GDP); the degree of cross-border financial interconnectedness (eg net non-resident loans to domestic banks etc); a general indicator of economic and financial development, such as the level of per-capita income (in US dollar PPP terms); measures of institutional quality (quality of regulation, rule of law, effectiveness of public administration, and the quality of banking supervision). Annex Table A.1 provides details about the data and sources.

We then carry out a specification search. We include variables that proxy for all these factors and look for a model that fits the data well, has plausible coefficients, and meets error distributional assumptions. We further reduce the model by dropping variables with the aim of improving predictive power, limiting the risk of over-parametrisation.

The outcome of this specification search is shown in Table 1. The estimates reported in this table are based on two sample periods. The first is the full sample. However, this sample does not include bank capital, which is a potentially relevant variable but available for a sufficient number of countries only after 2000. The second sample includes bank capital but, accordingly, covers only post-2000 crisis episodes.[18]

Across both samples, three variables are significant predictors of the (log) fiscal cost. The first is the ratio of *private non-financial credit* to GDP. Given the log-linear specification,[19] a 10 percentage point rise in this variable is associated with a point-estimate increase in the fiscal cost of 6–8%, all else equal.[20] The second is pre-crisis five-year *average credit growth*, which likely captures the deterioration of credit quality that occurs when credit rises more rapidly. A one percentage point increase in this variable corresponds to 2% and 5.2% higher fiscal cost in the full and the shorter samples, respectively. The third variable is *foreign exchange reserves* (as a share of GDP). A one percentage point increase in *foreign exchange reserves* (as a share of GDP) coincides with a lower cost of between 3.8% and 5.1% at the point estimate, depending on the sample.

17 We use a one-sided output gap to capture only information that is available to the authorities when estimating the costs, ie to approximate real-time information (see also Borio et al (2017) for the cyclical adjustments of fiscal positions to take the financial cycle into account).

18 We find that using capital to unweighted assets performs better, although the results are very similar when using the risk-weighted measure.

19 In (1) above, the coefficient $b = \partial \ln(c)/\partial x = (\partial c/c)/\partial x$. A change in one unit of the risk factor x is related to a change of b per cent in the cost c.

20 Measures of financial sector leverage, such as bank assets to GDP and the loan-to-deposit ratio, are statistically significant only when the private credit ratio is dropped.

Fiscal cost risk factors in benchmark regression

Table 1

Variables			Summary statistics			
	Full sample	Sample including bank capital	Full sample		Sample including bank capital	
	Coefficient	Coefficient	Mean	Standard deviation	Mean	Standard deviation
Dependent variable: Increase in public debt over a five-year window (in logs)			21.8	20.2	25.9	22.4
PNF credit to GDP (%)	0.008***	0.006***	68.2	53.5	87.0	59.1
	(0.001)	(0.002)				
PNF credit to GDP growth (five-year)	0.022**	0.052**	6.2	9.9	9.1	8.7
	(0.010)	(0.020)				
FX reserves (% of GDP)	–0.051***	–0.038**	8.5	7.9	9.6	9.7
	(0.012)	(0.017)				
Bank capital (% unweighted assets)		–0.103**			8.5	4.3
		(0.042)				
Government debt (% of GDP)	0.004***		54.9	65.4	40.0	29.7
	(0.001)					
Observations	67	31				
R-squared	0.47	0.73				

*** Indicates 1%. ** Indicates 5%. * Indicates 10%. OLS with robust standard errors (in parenthesis). The cost of a banking crisis is the increase in public debt between one period before the onset of the crisis and four periods after it. In the regression such cost is expressed in logs.

Bank capital (as a share of unweighted assets) is an important risk-mitigating factor. A 1 percentage point increase is associated with a reduction in the fiscal cost of about 10%, *all else equal.*[21] The significance of this variable and the fact that bank capital has played a crucial role in the financial reforms implemented in the aftermath of the GFC, suggests using the model with bank capital to predict potential future fiscal cost (see Section 3). A potential problem with this choice is the smaller number of observations. However, point estimates of the coefficients for the other risk factors remain precisely estimated. Moreover, once taking into account sampling uncertainty, their magnitudes are reasonably similar to those estimated in the full sample.

Table 1 also suggests that a weaker fiscal position pre-crisis, as reflected in the ratio of *government debt* to GDP, is a relevant risk factor. In the full sample, a 10 percentage point increase corresponds to an increase in the fiscal cost of about 4%. In the model with bank capital, the magnitude of the effect is almost identical but statistically insignificant at conventional confidence levels. (This is not shown in Table

[21] The quality of institutions is also potentially important. As a proxy for this variable, we included in the regressions an index from Abiad et al (2008) that measures the degree of adoption of risk-based capital adequacy principles, independence and legal power of the supervisory agency, institutional coverage of supervision, and effectiveness of examinations of banks. This measure is statistically significant, but also highly correlated with other variables. Furthermore, adding this variable leads to losing almost half of the observations in the estimation.

1 but more details are provided in Table 2 below).[22] For this reason, we drop this variable from model including bank capital.

Table 2 provides more information about the search process yielding the specification in Table 1.[23] Column (1) displays the set of risk factors that provides the best fit according to R^2 out of a much larger set of potentially relevant variables. This set comprises: the level and growth of credit to the private non-financial sector; foreign exchange reserves; government debt; the output gap; three-year average GDP growth; and the three-year depreciation of the real US dollar exchange rate. Other possibly relevant variables, including measures of financial sector leverage, cross-border financial flows (eg net non-residents loans to domestic banks), the current account and the other variables did not turn out to be statistically significant.

Columns 2–4 report what happens when we drop some or all of the last three variables in column 1, which are less stable across different specifications than the first four variables (not shown here). Among these models, the best one according to the Bayesian Information Criterion (BIC) is the one in column 4,[24] in which the log fiscal cost is well predicted by the level and growth of private non-financial credit, foreign exchange reserves and the level of government debt. The estimated coefficients of these variables change little across specifications, unlike the remaining variables.

Once we re-estimate the model including bank capital (column 5), this variable turns out to be statistically and economically significant. Despite the smaller number of observations, point estimates of the coefficients for private non-financial credit remain very similar to those estimated in the full sample and precisely estimated. This suggests that the variables are robust predictors of the fiscal cost. Moreover, credit growth and foreign exchange reserves remain statistically significant, although the former is economically more relevant and the latter somewhat less than in the full-sample specification. Estimated coefficients for government debt are almost identical but less precisely estimated – that is, statistically insignificant at conventional confidence levels. Dropping this variable improves both the AIC and BIC indicators (column 6).

[22] The primary balance is also relevant, but when added to a regression that includes government debt, both variables become statistically insignificant – it is hard to tell their effects apart.

[23] It reports estimates for both advanced and EM economies. Unfortunately, we do not have enough observations to obtain statistically robust results when separating advanced economies from EMEs.

[24] The Bayesian Information Criterion (BIC) penalises complexity more than the Akaike Information Criterion (AIC) (also displayed in Table 1), thus reducing the risk of overparameterisation.

Risk factors affecting the log cost of banking crisis						Table 2
	(1)	(2)	(3)	(4)	(5)	(6)
PNF Credit to GDP	0.008***	0.008***	0.008***	0.008***	0.007***	0.006***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
PNF Credit to GDP growth (five-year)	0.025**	0.022**	0.024**	0.022**	0.056**	0.052**
	(0.011)	(0.011)	(0.010)	(0.010)	(0.021)	(0.020)
FX reserves (% of GDP)	–0.054***	–0.058***	–0.050***	–0.051***	–0.036**	–0.038**
	(0.013)	(0.012)	(0.012)	(0.012)	(0.017)	(0.017)
Government debt (% of GDP)	0.006***	0.005***	0.004***	0.004***	0.004	
	(0.001)	(0.001)	(0.001)	(0.001)	(0.004)	
Output gap (one sided)	–0.114***	–0.110***	–0.072**			
	(0.037)	(0.031)	(0.028)			
Average GDP growth (three-year)	0.078*	0.061				
	(0.037)	(0.031)				
Real dollar exchange rate (%ch 3yr)	0.010**					
	(0.004)					
Bank capital (% unweighted assets)					–0.092**	–0.103**
					(0.041)	(0.042)
Observations	67	67	67	67	31	31
R-squared	0.53	0.51	0.49	0.47	0.74	0.73
AIC	147.0	148.5	149.2	149.9	58.4	57.7
BIC	164.7	163.9	162.4	160.9	67.0	64.8

*** Indicates 1%. ** Indicates 5%. * Indicates 10%. OLS with robust standard errors (in parenthesis). The cost of a banking crisis is the increase in public debt between one period before the onset of the crisis and four periods after it. In the regression, this cost is expressed in logs.

We also performed two further checks. First, we tried to estimate a model with interaction dummies denoting advanced and non-advanced economies (the latter being emerging market economies (EMEs) and less developed economies). It turns out that the interactions were generally not statistically significant. Only the differences in coefficient estimates for the output gap, average GDP growth and currency depreciation appeared statistically significant. Yet, a model with interactive dummies on these three variables turns out to underperform in terms of the AIC criterion the best performing one in Table 1. Second, we checked whether the exchange rate regime matters. To this end, we created a dummy indicating peg or non-peg regimes based on the classification of Iltzetzki et al (2017). Specifically, we classify as "pegs" regimes classified as 1 and 2 in of Iltzetzki et al (2017) and as "non-pegs" all the others. We did not find statistically significant differences in the coefficients.

The results in Table 2 could be affected by a sample selection bias. As noted in Section 1, a small number of observations in our initial sample correspond to episodes in which debt declined after a banking crisis owing to sovereign defaults, debt restructuring or other reasons. To rule out this possibility, we estimated the regression model with the Heckman two-step correction procedure and found that the results are robust (Annex B).

3. Predicting the fiscal costs

We next estimate the ex ante distribution of fiscal costs employing the same method as in Adrian et al (2019). We start by estimating quantile regressions of the fiscal cost. We then smooth the quantile distribution by interpolating the predicted quantiles using a skewed t-distribution. This method is flexible enough to allow the distributions to change over time. Formally, we first estimate quantile regressions of the log fiscal cost y_{it+h} on the risk factors x_{it} where the slope β_τ for the quantile τ is given by

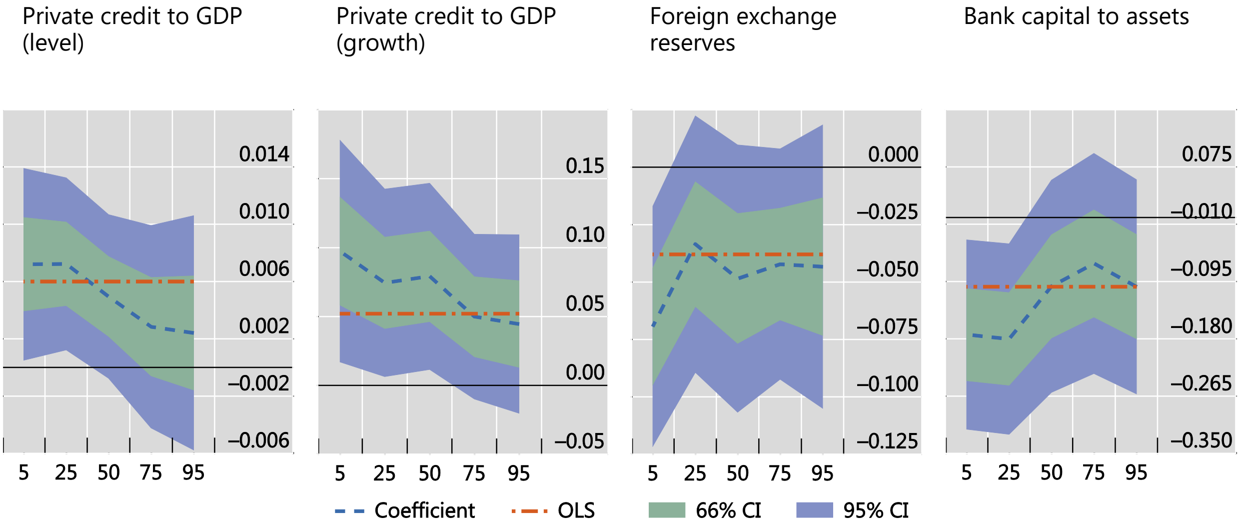
$$\widehat{\beta_\tau} = argmin \sum_{t=1}^{T-h} \big(\tau \cdot 1_{(y_{it+h} \geq x_{it}\beta)} \, |y_{it+h} - x_{it}\beta| + (1-\tau) \cdot 1_{(y_{it+h} < x_{it}\beta)} \, |y_{it+h} - x_{it}\beta|\big) \tag{2}$$

in which 1 is an indicator function. We use the model that includes bank capital (column 7) because it is the one that turns out to give tighter distribution for all countries: that is, although the expected values that each model delivers are very similar, the tails of the distribution are more precisely estimated when bank capital is included.[25]

We estimate coefficients for four quantiles: 5th, 25th, 75th and 95th. As will be clearer below, those few points increase the flexibility in estimating the shape and location parameters of the distribution. For each of them, Figure 2 shows the estimated coefficients along with their 66% and 95% confidence intervals. For comparison, Figure 2 also shows the corresponding OLS estimates, which remain around the value of the coefficients in the quantile regression.

25 We apply the quantile regression on the baseline model and not on the selection model in Annex B for two reasons. First, the coefficients for the risk factors are very similar. Second, it is not straightforward to incorporate the selection equation into the quantile procedure and into the forecast of the distribution.

Coefficients of the risk factors, by quantile[1] Figure 2



Note: The blue line shows the coefficients for each factor of quantile regression for the 5th to the 95th quantile where the dependent variable is the cost of banking crisis. The red line shows the coefficient of the OLS regression version. Confidence intervals at the 66% and 95%.

Source: Author's calculations.

Once we have estimated the quantile regressions (2), we can compute the predicted quantile values of the fiscal cost for each set of observations of the risk factors x_{it} corresponding to a pair (country, year) as:

$$\hat{Q}_{y_{it+h}|x_t}(\tau|x_{it}) = x_{it}\hat{\beta}_\tau \tag{3}$$

Each predicted quantile in (3) corresponds to a point in the *cdf* F(.) of the log fiscal cost. These points can be interpolated to obtain a smooth distribution with which we can carry out a VaR-type analysis. Smoothing is also needed because quantile predictions are often noisy, given the error associated with the estimation. In our case, we interpolate semiparametrically the predicted quantiles using the skewed *t*-distribution to take advantage of its overall shape flexibility. The distribution is described by the following function:

$$f(y;\mu,\sigma,\alpha,v) = \frac{2}{\sigma}t\left(\frac{y-\mu}{\sigma};v\right)T\left(\alpha\frac{y-\mu}{\sigma}\sqrt{\frac{v+1}{v+\frac{y-\mu}{\sigma}}};v+1\right) \tag{4}$$

where *t(.)* and *T(.)* are the *pdf* and the *cdf* of the student *t*-distribution. The four parameters that control the location (μ), scale (σ), fatness (v) and shape (α) of the distribution can be computed for each set of observations corresponding to a pair (country; year) to minimise the distance between the estimated quantiles and those implied by the distribution, ie:

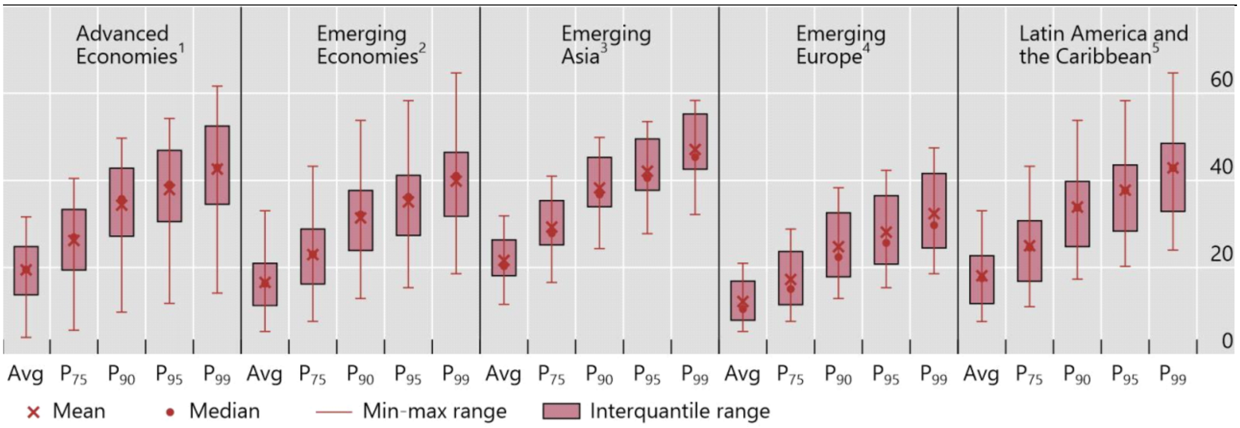
$$(\hat{\mu}_{it+h},\hat{\sigma}_{it+h},\hat{\alpha}_{it+h},\hat{v}_{it+h}) = argmin\sum_{\tau}\left(\hat{Q}_{y_{it+h}|x_t}(\tau|x_{it}) - F^{-1}(\tau;\mu,\sigma,\alpha,v)\right)^2 \tag{5}$$

For each pair (country, year) we thus obtain a distinct distribution with which we can compute the mean, the 75th, the 90th, the 95th and the 99th percentiles.

Estimated fiscal cost of banking crisis should a crisis occur, as of 2018

Percentage of GDP

Figure 3



[1] Advanced economies: Australia, Austria, Belgium, Croatia, Cyprus, Denmark Finland, Greece, Ireland, Luxembourg, Malta, the Netherlands, Portugal, Slovakia, Spain and Switzerland. [2] Emerging economies: Algeria, Argentina, Belarus, Bolivia, Bosnia and Herzegovina, Brazil, Bulgaria, Chile, China, Colombia, the Czech Republic, the Dominican Republic, Ecuador, El Salvador, Estonia, India, Indonesia, Kazakhstan, Latvia, Lithuania, Malaysia, Mexico, Moldovia, Nicaragua, Nigeria, North Macedonia, Pakistan, Panama, Paraguay, Peru, the Philippines, Poland, Russia, South Africa, Sri Lanka, Thailand, Turkey, Ukraine, Uruguay, Vietnam and Zambia. [3] Emerging Asia: China, India, Indonesia, Malaysia, the Philippines, Sri Lanka, Thailand and Vietnam. [4] Emerging Europe: Belarus, Bosnia and Herzegovina, Bulgaria, the Czech Republic, Lithuania, Latvia, Moldovia, North Macedonia, Poland, Russia, Turkey and Ukraine. [5] Latin America and the Caribbean: Argentina, Brazil, Bolivia, Chile, Colombia, the Dominican Republic, Ecuador, El Salvador, Mexico, Nicaragua, Panama, Paraguay, Peru and Uruguay.

Note: The fiscal cost of a banking crisis is the predicted increase in public debt over a five-year period conditional on a crisis occurring and the observation of the risk factors at the end of 2017. Avg indicates the mean fiscal cost given a crisis; P75 P90, P95 and P99 are the 90th, 95th and 99th quantiles, respectively. The graph shows how the estimates of the average and the quantiles of the fiscal cost are distributed across countries.

Source: Authors' calculations.

Figure 3 presents the estimated distribution of fiscal cost (increase in public debt over a five-year window) as of 2018 for different groups of countries. The expected average fiscal cost is within a range of 5–30 % of GDP for advanced economies, with a cross-country average of 20%. The range is similar for EMEs and less developed economies, although the cross-country average is lower. At the tail of the distribution, the fiscal losses can be substantial. On average across advanced economies, the 95th quantile is about 38% of GDP and the 99th quantile exceeds 40%. For some countries, the 95th and 99th percentile losses can be as high as roughly 50% and 60% of GDP. For EMEs, the cross-country averages of the highest quantiles are generally close but lower than those for advanced economies.

Taking these estimates at face value, a banking crisis would reduce fiscal space significantly. At the same time, if we consider specific measures such as those presented in Fournier and Fall (2015), the available space appears sufficient for most of the advanced economies in our sample. That said, the measures of fiscal space in Fournier and Fall (2015) and similar studies are deterministic and hence probably overstate the true amount of space available significantly. In the case of EMEs, considering the calculations of Ganiko et al (2016), which do account for various sources of uncertainty, fiscal sustainability in a number of countries looks vulnerable in the case of serious financial stress.

Some important qualifications are in order. First, our estimates include only one element of the wide-ranging post-GFC financial reforms – higher bank capital relative to assets. Bank capital alone can have a sizeable effect on fiscal costs. As an illustration, we run a counterfactual exercise in which we estimate what the fiscal cost of the crisis in 2008 would have been, had the level of bank capital in 2007 been the same as in 2017, all else equal. Using the model for the mean, the fiscal cost would have fallen by some 7 percentage points of GDP on average across countries; using the model for the 95th percentile, the fall would have been equal to 12 percentage points, all else equal. In the case of some countries, notably Ireland, the reduction in cost would have been much higher. More generally, the estimated reductions would likely have been considerably larger had we considered also the impact of the other post-GFC reforms, including other improvements in the quality and robustness of bank capital, the introduction of liquidity standards, improvements in the resolution of systemically important institutions so as to avoid the use of taxpayers' money (ending "too-big-to-fail") and the implementation of macroprudential frameworks (eg BIS (2018)).[26]

Another qualification is of a technical nature. Although we employ a semiparametric method to fit a flexible distribution of potential costs, distributional assumptions matter. For example, assuming a symmetric t-distribution (which is a special case of the skewed t distribution used here) would generally lead to higher fiscal buffer estimates because of larger standard errors. More generally, while we are confident of the quality of our estimates, any estimate is inevitably subject to uncertainty, reflecting the method used, data availability and the necessarily simplified framework, including the assumption that the parameter estimates are uniform across countries. This should be taken into account when setting policy (see below).

Conclusion

Banking crises can result in large increases in public sector debt – a measure of their fiscal cost. In this paper we find that a relatively small number of factors can explain a significant share of the variation of fiscal costs in past banking crises. Furthermore, based on these findings, we propose a method to estimate the distribution of the potential future fiscal loss given a banking crisis. This helps assess the adequacy of fiscal buffers. The measure is akin to the value-at-risk routinely used in risk management and more recently employed in macroeconomic analysis in the form of GDP-at-risk (Adrian et al (2019)). While our methodology is intended only as one of the inputs in any decision regarding the adequacy of fiscal buffers, it provides a sense of the orders of magnitude involved. So measured, the magnitude is sizeable.

Our findings have a number of policy implications. First, they stress the importance of maintaining or rebuilding fiscal buffers in good times, when it may in fact be tempting to relax fiscal policy. This is especially the case if good times go hand in hand with outsize cumulative credit growth in the private sector. A large body of empirical evidence indicates that this is a symptom of the build-up of vulnerabilities that can herald financial crises. Our results complement that work by suggesting that

[26] It is reasonable to believe that the reforms, notably the implementation of macroprudential frameworks, have also reduced the impact of credit growth on the cost of banking crises.

it is also a useful indicator of the crises' fiscal costs. Second, our findings point to the importance of prudential regulation in limiting those costs. In our study we can only use the level of bank capital in relation to assets. But other measures no doubt matter in this context, not least those designed to ensure orderly resolution and minimise any resort to tax payers' money as well as the implementation of macroprudential frameworks. Finally, by exploring one key link between fiscal policy and financial instability, our work takes a further step in developing a holistic macro-financial stability framework. In such a framework, micro- and macroprudential policies work alongside monetary, fiscal and even structural policies to ensure sustainable financial and macroeconomic stability (BIS (2018)).[27]

[27] Another step is estimating (real-time) cyclically adjusted fiscal positions by taking the financial cycle explicitly into account, as in Borio et al (2017). Such adjustments can be sizeable, as financial booms greatly flatter fiscal accounts.

Annex A: Data sources

Description of the variables Table A1

Variable	Description	Source
PNF credit to GDP	Private credit by deposit money bank and other financial institutions to GDP (%).	Data up to 2016: T Beck, A Demirgüç-Kunt and R Levine, "A New Database on Financial Development and Structure", *World Bank Economic Review*, no 14, 2000, July 2018 release update, pp 597–605. Data for 2017: IMF, IFS. Authors' calculations.
FX reserves (% of GDP)	Total foreign exchange reserves excluding gold as a percentage of GDP.	IMF, WEO (both variables: reserves and GDP).
Government debt (% GDP)	Public debt to GDP ratio.	Data up to 2015: IMF, Historical Public Debt Database. Debt to GDP Ratio Data for 2016-2017: IMF, WEO. General Government Debt
Primary fiscal balance	% General government primary balance as a GDP percentage.	IMF, WEO.
Per-capita income (PPP USD)	Real GDP per capita	IMF, WEO.
Bank capital (% unweighted assets)	Bank capital to assets ratio (%)	World Bank, World Development Indicators.
Real dollar exchange rate	Bilateral US dollar real exchange rate constructed as the nominal end-of period exchange rate against the US dollar times the US GDP deflator and divided by the domestic GDP deflator.	Nominal exchange rate from the Global Financial Data. GDP deflator from World bank, World Development Indicators.
Nominal effective exchange rate	Nominal Effective Exchange rate, nominal (Narrow index: 26 cty) (BIS),D, A-Avg.	BIS
Output	Nominal GDP.	IMF, WEO.
Indicator for exchange rate regimes	Dummy for peg/non peg regime	Iltzetzki et al (2017)
Current account balance	Current account balance	IMF, WEO.
Quality of regulation	Reflects perceptions of the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development.	Worldwide governance indicators (WGI) project, World Bank
Rule of law	Reflects perceptions of the extent to which agents have confidence in and abide by the rules of society, and in particular the quality of contract enforcement, property rights, the police, and the courts, as well as the likelihood of crime and violence.	Worldwide governance indicators (WGI) project, World Bank
Effectiveness of public administration	Reflects perceptions of the quality of public services, the quality of the civil service and the degree of its independence from political pressures, the quality of policy formulation and implementation, and the credibility of the government's commitment to such policies.	Worldwide governance indicators (WGI) project, World Bank
Quality of banking supervision	Index for the WB: based on several measures of banking activity regulatory variables, capital regulatory variables, official supervisory action variables, and deposit insurance scheme variables	World Bank Surveys on Bank Regulation and Financial Reform Database (IMF)

Annex B: Heckman two-step correction for selection bias

In this annex we use Heckman's (1979) two-step procedure to check for selection bias. In the first step, we estimate a probability model to account for the observations that we do not observe. In the second step, we estimate the risk factors that affect debt increases.[28]

Heckman two-step selection model for the log fiscal cost

Table B1

	Full sample model		Bank capital model	
	Log(C)	Selection	Log(C)	Selection
PNF credit to GDP	0.008***	-0.001	0.005***	-0.006
	(0.002)	(0.004)	(0.001)	(0.004)
PNF credit to GDP growth (five-year)	0.024**	-0.006	0.062***	0.159**
	(0.011)	(0.021)	(0.015)	(0.071)
FX reserves (% of GDP)	–0.048***	-0.020	–0.044***	-0.034
	(0.012)	(0.024)	(0.014)	(0.052)
Government debt (% of GDP)	0.005***	0.000		
	(0.002)	(0.002)		
Average GDP growth (three-year)		0.117*		
		(0.063)		
Real dollar exchange rate (%ch 3 yr)		0.012		0.289
		(0.10)		(0.205)
Bank capital (% unweighted assets)			–0.108***	-0.122
			(0.030)	(0.742)
Lambda	-0.47		0.55	
std	(0.37)		(0.09)	
Chi-squared	52.77		78.03	
Uncensored observations	67		31	
Censored observations	13		6	

*** Indicates 1%. ** Indicates 5%. * Indicates 10%. OLS with robust standard errors (in parenthesis).

Formally, we model the first stage as $P(Debt\ increase = 1|Z) = \varphi(Z\delta)$, where Z are the factors that explain the probability of observing an increase in debt and δ are parameters. If debt increases, $C = \log(c)$ is observable, otherwise we do not observe it. We estimate the risk factors associated with debt increases as $C = X\beta + \omega$. We then combine both equations and estimate the risk factors given that we observe an increase in debt:

$$E(C|X, debt\ increase = 1) = X\beta + E(\omega|X, debt\ increase = 1)$$

Under the assumption that the error terms are jointly normal, this equation is

$$E(C|X, debt\ increase = 1) = X\beta + \rho\sigma\omega\lambda(Z\delta)$$

28 The analysis of the factors that determine the occurrence of banking crisis is beyond the scope of this paper. This annex is only intended to show that the selection bias in our sample does not affect our conclusions about the risk factors influencing the cost of a crisis.

where ρ is the correlation of the error terms in the equation, σw is the standard deviation of w and λ(Zδ) is the inverse mills ratio evaluated at Zδ.

Annex Table B.1 shows that the estimates in Table 1 are robust to selection effects. We choose a selection equation that gives a better overall fit for the model. The coefficients in the estimating equation are very similar to the coefficients in Table 1, although some lose statistical significance.

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