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Monetary facts revisited*

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Abstract

This paper uses a cross-country database covering 46 economies over the post-war period to revisit two key monetary facts: (i) the long-run link between money growth and inflation and (ii) the link between credit growth and financial crises. The analysis reveals that the former has weakened over time, while the latter has become stronger. Moreover, the money-inflation nexus has been stronger in emerging market economies than in advanced economies, while it is the other way round for the link between credit growth and financial crises. These results suggest that there is an inverse relationship between the two monetary facts. The money-inflation link is weaker in regimes characterised by low inflation and highly liberalised financial systems, while the reverse holds true for the credit-crisis nexus.

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1 Introduction

In the decades leading up to the Great Financial Crisis, money and credit aggregates played an increasingly peripheral role in monetary theory and policy. In mainstream macro models, money and credit did not matter for macroeconomic outcomes.[1] And in monetary policy frameworks, money and credit were generally not amongst the key indicators considered.[2] This marginalisation of money and credit has been a subject of debate, both before and, more intensely, after the crisis.

Before the Great Financial Crisis, the debate focused on the role of money in monetary theory and policy. Specifically, monetary policy frameworks and monetary macro models were criticised by some for disregarding the long-run one-to-one link between money growth and inflation that is predicted by the quantity theory of money. This long-run link has been a core monetary fact of the post-war period, strongly supported by the empirical evidence. Specifically, numerous empirical studies have shown that the long-run averages of money growth and inflation are proportionally correlated across countries (e.g. McCandless and Weber (1995), Lucas (1996), Vogel (1974), Lothian (1985) and Dwyer and Hafer (1988, 1999)).[3] Indeed, Figure 1 (left-hand panel) shows that this evidence can also be reproduced from up-to-date data. The averages of money growth (in excess of real GDP growth) and average inflation in a group of 46 economies are correlated on a one-to-one basis over the post-war period.

The Great Financial Crisis has lent new impetus to the debate about the role of quantitative aggregates in monetary analysis and modelling, but with a shift in focus. The financial crisis was preceded by a credit boom, which turned into a bust when the crisis broke out, a monetary fact that has generally held over the post-war period (Figure 1, right-hand panel).

[1] The key feature of pre-crisis mainstream New Keynesian models was pricing frictions, while financial factors were essentially absent with the exception of a short-term interest rate controlled by the central bank. In its simplest form, the baseline New Keynesian model could be boiled down to a three-equation system featuring inflation, the output gap and the short-term interest rate (see Woodford (2008)).

[2] That said, over much of the post-war period, money and credit aggregates figured prominently in many countries' monetary policy frameworks. In the 1950s and 1960s, credit played an important role when many central banks had put in place restrictions on interest rates and balance sheet quantities through which they implemented credit allocation and demand management policies. The 1970s then saw the emergence and spread of monetary targeting and of an increasing focus on money at the expense of credit in policymaking. This shift was driven by the ascent of the monetarist paradigm, which emphasised the implications of the quantity theory of money for macroeconomic outcomes, in particular the long-run association between money growth and inflation. Since the mid-1980s, there was a gradual move away from monetary targets and towards more directly inflation-centred regimes. See Borio and Lowe (2004) and BIS (2007) for a more detailed discussion of the evolution of the role of credit and money in monetary policy frameworks over the post-war period.

[3] Another strand of the literature demonstrated the existence of a long-run link between money growth and inflation based on long runs of time series data for individual countries (see e.g. Lucas (1980), Assenmacher-Wesche and Gerlach (2007)).

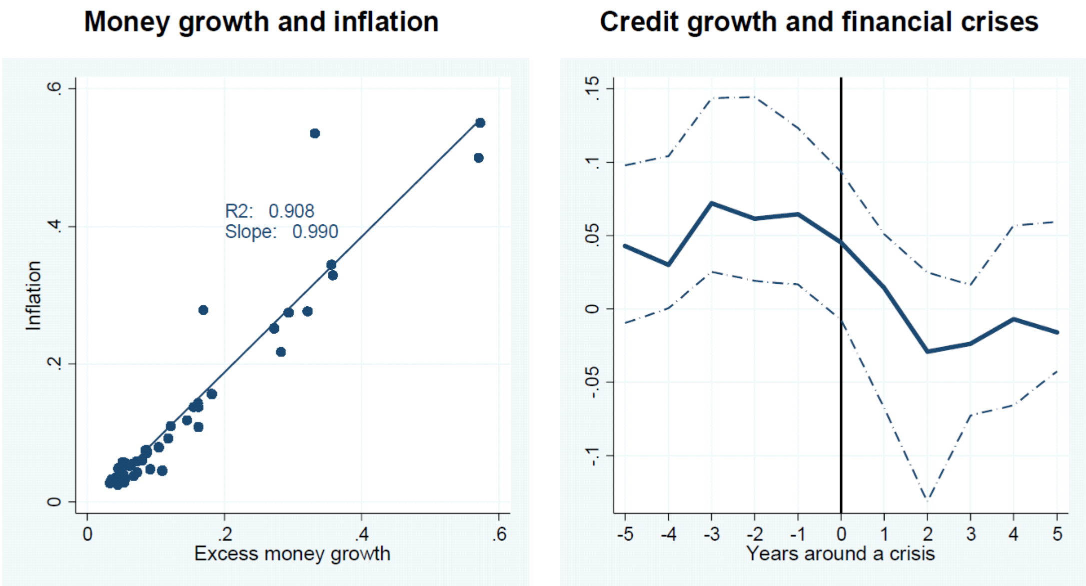

Figure 1. **Two monetary facts.** The left-hand panel shows pairs of country averages of the log change in the GDP deflator and the log change in broad money over real GDP. The right-hand panel shows the median and the interquartile range acrosss countries of the log change in the credit-to-GDP ratio around financial crisis events. The calculations are based on data for 46 countries over the period 1950-2011 (see Annex-Table A1). Sources: BIS, IMF, Global Financial Data, St. Louis Fed FRED database, national sources, authors' calculations.

This observation has reinforced calls that credit aggregates ought to be given greater attention in monetary analysis in order to better identify risks to financial stability and ultimately to long-run price stability. Borio and Lowe (2002a, b, 2004) first emphasised and provided formal evidence on the link between credit and financial instability. They showed, based on cross-country empirical evidence, that persistent growth of credit above its long-term trend indicates a growing risk of a systemic financial crisis.[4] In a similar vein, Eichengreen and Mitchener (2003) characterised the Great Depression as a credit boom gone bust. Recently, Schularick and Taylor (2012) documented the significant predictive ability of real credit growth for future financial crises for a historical panel of advanced economies over the period 1870-2008. Borio and Drehmann (2009), Gourinchas and Obstfeld (2012) and Drehmann and Juselius (2013) show that domestic credit-based indicators have been reliable leading indicators of financial crises in both advanced and emerging market economies over the post-Bretton Woods period.[5]

[4]The early literature on the leading indicators of banking crises had identified credit growth as one important leading indicator amongst others (e.g Demirgüc-Kunt and Detriargache (1998), Kaminsky and Reinhart (1999)).

[5]Borio and Drehmann (2009) and Drehmann and Juselius (2013) suggest that the credit gap, the deviation of credit from a long-run trend, is a reliable indicator of financial distress. Gourinchas and Obstfeld (2012) find that, in addition to credit growth, real currency appreciation is also an important indicator. Another

Monetary facts might, however, change over time and differ between countries due to changes and cross-country differences in the monetary and financial regimes. Two regime changes that stand out over the post-war period are the significant global disinflation and financial liberalisation trends since the mid-1980s (Figure 2).[6] The global median inflation rate dropped from 13% to 7% in the mid-1980s, and then below 5% in the mid-1990s (left-hand panel). Financial liberalisation, measured by the quantitative indicator of Abiad et al. (2010), which ranges between 0 (full repression) and 1 (full liberalisation),[7] accelerated globally in the early-1980s (right-hand panel). Another liberalisation wave followed in the early 1990s.

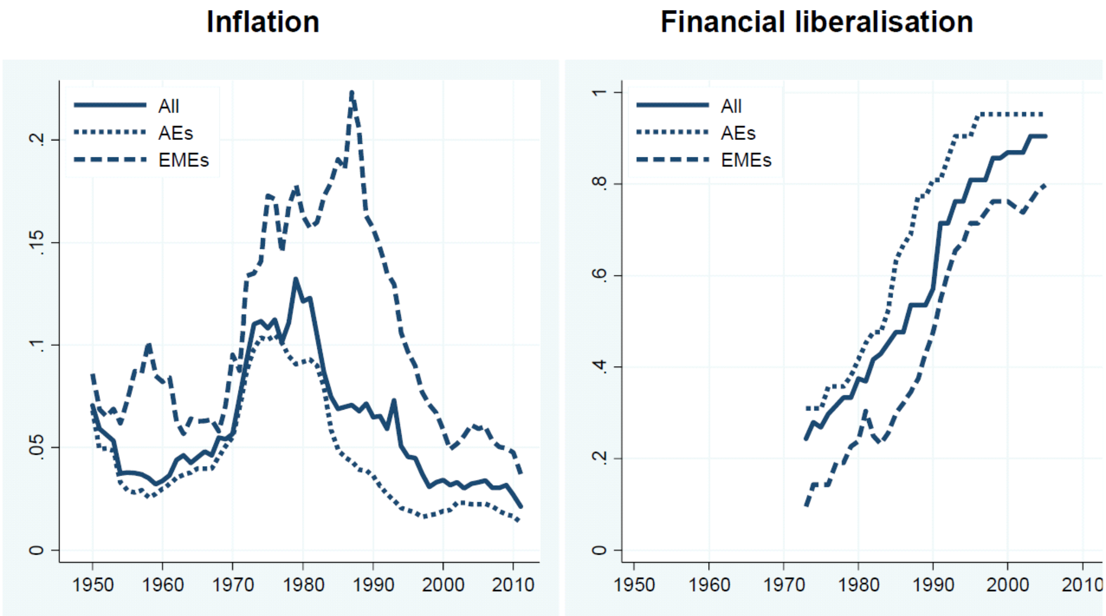

Figure 2. **Disinflation and financial liberalisation**. The charts display cross-country medians of GDP deflator inflation and of the financial liberalisation index of Abiad et al. (2010). The latter ranges from 0 (full repression) to 1 (full liberalisation). Sources: Abiad et al. (2010), BIS, IMF, Global Financial Data, St. Louis Fed, national sources, authors' calculations.

The graph also reveals that there are important differences between advanced and emerging market economies. Specifically, inflation has been on average higher in emerging market economies (EMEs) than in advanced economies (AEs), and widespread disinflation set in

strand of literature has established a significant empirical link between money and credit growth and asset price dynamics (e.g. Borio et al (1994), Detken and Smets (2004), Adalid and Detken (2007), Goodhart and Hofmann (2008), Alessi and Detken (2009)).

[6] The two trends are probably not entirely unrelated to each other. Their coincidence may reflect the retreat of policies using a combination of high inflation and financial repression to liquidate government debt (Reinhart and Sbrancia (2011)). It may also reflect that financial deregulation strengthening the transmission of monetary policy was in many countries an important element of disinflation strategies (Pagoulatos (2003)).

[7] The indicator of Abiad et al (2010) is based on a grading of seven dimensions of financial sector policy, with the grading ranging from zero (no liberalisation) to 1 (full liberalisation) for every year between 1973 and 2005. A higher value of the indicator thus reflects a more liberal financial system.

later, with median inflation rates dropping sharply only since the early 1990s. Also financial systems are on average less liberalised in EMEs than in advanced economies.

Financial liberalisation and disinflation may weaken the empirical link between money growth and inflation in two main ways. First, velocity shifts due to financial innovation or because of a change in the regime rate of inflation may drive a wedge between money growth and inflation, thus reducing the reliability of money growth as an indicator of inflation (Lucas (1988), Reynard (2006), McCallum and Nelson (2011)). Second, in an environment of low and stable inflation, the link between money growth and inflation may become blurred as velocity shocks play a more dominant role, obscuring the signal from money growth (Estrella and Mishkin (1997), De Grauwe and Polan (2005).

The evidence supports the notion that these forces are at work. For instance, Friedman and Kuttner (1992) and Estrella and Mishkin (1997) showed for the United States and Germany that the relationship between money growth and inflation has vanished since the early 1980s. De Grauwe and Polan (2005), Teles and Uhlig (2013) and Teles et al. (2015) have presented evidence suggesting that the long-run cross-country link between average inflation and average money growth weakens or entirely disappears in low inflation environments. Similarly, Bordo and Filardo (2007), Benati (2009) and Sargent and Surico (2010) present historical time series evidence showing that the long-run association between money growth and inflation is weaker when inflation is low.

McCallum and Nelson (2011) have recently however questioned this evidence. They argue that the widespread approach of analysing the money-inflation nexus based on averages, either cross-country averages in panel studies or moving averages (or other types of low-frequency filters) in time series studies ignores the dynamic lead-lag relationship between the two variables. Specifically, they argue that money growth tends to lead inflation and that failing to take this into account will lead to an underestimation of the strength of the link, a point that they substantiate by empirical evidence from time series analysis for G7 countries.

The potential impact of financial liberalisation and disinflation trends on the association between credit and financial crisis is also twofold, but working in the direction of strengthening the link. Bordo et al. (2001) have documented that the frequency of financial crises has increased considerably in the post-Bretton Woods period. The coincidence with the trend towards more liberal financial systems over the same period suggests that there might be a link between financial liberalisation, credit boom-bust cycles and financial crises. Indeed, the link between financial liberalisations and subsequent credit booms and busts is well documented in the literature (e.g. Goodhart et al. (2004)).

At the same time, the advent of low and stable inflation regimes, buttressed by central bank anti-inflation credibility and the disinflationary forces of globalisation, may have changed the way unsustainable economic expansions manifest themselves. Instead of showing up first and foremost in rising inflation, they may now become visible primarily in unsustainable increases in credit and asset prices that then usher in a financial crisis (Borio et al. (2003), Borio and Lowe (2004), Borio (2005), Borio (2012)). The reasoning is the following. When inflation is held down by anchored inflation expectations and global disinflationary forces, the build-up of economic and financial imbalances will have a lesser impact on short-term inflation. And when monetary policy is, in turn, focused on short-term inflation developments as prescribed in standard inflation-targeting-type frameworks, it will unwittingly accommodate the build-up of these imbalances and ultimately of future risks to financial stability.

There is no direct evidence available that the association between credit and financial crisis might have become stronger in the wake of financial liberalisation and disinflation. Schularick and Taylor (2012) perform a sub-sample analysis of their historical dataset, but they distinguish between a pre- and post-war sample and do not consider any potential further changes over the post-war period. Other strands of the literature do, however, indicate changes in the way financial shocks have affected asset prices since the mid-1980s, when financial systems were liberalised and price stability was established. For example, Goodhart and Hofmann (2008) show that credit shocks have had a considerably larger impact on house prices in OECD countries since the mid-1980s. Eickmeier and Hofmann (2013) and, more recently, Hofmann and Peersman (2016) demonstrate that monetary policy shocks have a significantly larger effect on credit and house prices in the United States over the same period.

The analysis in this paper aims to more systematically address the question whether the significant changes in the financial and monetary regime in the post-war period have affected monetary facts. To this end, we revisit the money-inflation and credit-crisis nexus based on a dataset covering 46 major advanced and emerging market economies over the period 1950-2011, using annual data. Specifically, we estimate the link between money growth and inflation and between credit growth and financial crises using panel econometric techniques. Within this econometric framework, we assess potential changes over time based on sub-sample analysis, re-estimating the two monetary relationships over the post-1984 and post-1994 sample period. Moreover, we also assess differences in the two monetary facts between advanced and emerging market economies against the background of the significant differences between the two country groups in terms of inflation performance and the degree of financial liberalisation.

In a nutshell, the results suggest that the money-inflation link has indeed weakened over time, while the credit growth-financial crisis link has become stronger. We further find that the association between money growth and inflation is stronger in emerging market economies (EMEs) than in advanced economies (AEs), while we find the opposite for the nexus between credit growth and financial crises. This suggests that there is an inverse relationship between the two monetary facts. The money-inflation link is weaker in regimes characterised by low inflation and highly liberalised financial systems, while the reverse holds true for the link between credit growth and financial crisis.

The paper further contributes to the literature in two other specific ways. First, we address the McCallum and Nelson (2011) criticism of the existing literature on the money-inflation nexus. We use the Pooled Mean Group (PMG) estimator proposed by Pesaran et al. (1999) as an effective way of addressing the caveats they raised in a multi-country panel framework.

Second, we explore potential changes over time in the credit-crisis nexus. Ongoing attempts to integrate the link in monetary policy frameworks often rely on relationships estimated on long historical data.[8] We assess whether the link has changed due to recent disinflation or financial liberalisation.

The remainder of the paper is structured as follows. Section 2 presents the data used in the empirical analysis. Section 3 outlines the econometric approach. Section 4 presents the main empirical findings. Section 5 performs a number of robustness checks. Section 6 concludes.

2 Data

The analysis is based on annual data for 46 major advanced and emerging market economies over the period 1950-2011. The country and time series sample covered by the analysis is reported in Annex-Table A1.

The data series used in the main part of the analysis are real and nominal GDP, the GDP deflator, broad money, bank credit to the private non-financial sector and an indicator of systemic financial crisis events. The data for GDP and the GDP deflator are from the IMF International Financial Statistics (IFS), Global Financial Data and national sources. Broad monetary aggregates are either M2 or M3 from the FRED database of the St. Louis Fed, complemented by the broad monetary aggregate from the IMF IFS (money plus quasi-

[8]See e.g. Sveriges Riksbank (2013) for considerations on how to integrate concerns about increasing household indebtedness in Sweden into the monetary policy decision-making framework of the Swedish Riksbank. These considerations feature two main parameters: (i) how monetary policy impacts household indebtedness, and (ii) how a change in household indebtedness (or credit expansion) impacts the probability of a financial crisis. For a critical discussion, see Svensson (2014).

money, series codes 34 and 35). For the euro area countries, national contributions to M3 are used back to 1970 and broad money from either national sources or the IFS. The bank credit series are from the BIS database on credit to the private sector, as described in detail in Dembiermont et al. (2013).[9] It is defined as credit extended by domestic banks to non-financial private sectors of the economy. The BIS data are complemented by compatible bank credit series from the IFS (series code 32.d) and from national sources.

The indicator variable for banking crisis events is constructed as a dummy variable that equals one when a crisis began in a country in period t, and zero otherwise. We focus on episodes of systemic financial crisis rather than more widely defined episodes of financial distress affecting only part of the financial system.[10] In this vein, we use the post-war financial crisis dates of Schularick and Taylor (2012) for the 14 advanced economies covered in their analysis. For the other countries, we use the crisis dates from Laeven and Valencia (2008, 2012) for period 1970-2011. For the pre-1970 period, we base the dating on Bordo et al. (2001) and Reinhart and Rogoff (2009).[11]

The analysis is performed on log differences of the data (except, of course, for the financial crisis indicator variable).[12] Table 1 provides a few summary descriptive statistics for the log change in the GDP deflator, real GDP, broad money and domestic bank credit. The table also reports for each series the results of the Im-Pesaran-Smith (IPS) panel unit root test (Im et al. (2003)). The statistics are reported for the full group of countries and for the group of advanced and emerging market economies separately.

The table reveals that there is considerable cross-country variation in inflation rates, with a significant number of high inflation outliers reflected in the large difference between the mean and median inflation rates (12.5% vs 5.4%) and a very high standard deviation of 29.4 percentage points. This pattern is driven by the EMEs, which have a considerably higher mean and median inflation rate (19.8% and 8.4% respectively) and a considerably larger standard deviation (39.6 percentage points) than the AEs (4.9%, 3.5% and 4.5 percentage points respectively).[13]

The variation in real GDP growth rates is much smaller, with very similar mean and

[9] The database is publicly available at http://www.bis.org/statistics/credtopriv.htm.

[10] See Bordo and Meissner (2016) for a comparison of financial crisis dates provided by different studies in the literature. They find that there are sometimes significant differences, reflecting differences in the stringency of the definition of a crisis. For instance, the definition in Laeven and Valencia (2008, 2012) is more stringent that that in Reinhart and Rogoff (2009), thus identifying fewer financial crisis events.

[11] Laeven and Valencia (2012) provide crisis codings only for the period 1970-2011.

[12] This is not a nuisance since the quantity theory implies a long-run relationship between the log difference of the price level and the log difference of money, not between the percentage rates of change. This point has been emphasised by Frain (2004).

[13] The high inflation outliers reflect hyperinflations that occurred in a number of countries (Argentina, Bolivia, Brazil, Chile, Indonesia, Israel, Mexico, Peru, Russia, Turkey, Uruguay and Venezuela).

	Mean	Median	S.D.	Unit root	N	NxT
Dlog price level						
All	0.125	0.054	0.294	-12.04	46	2,647
AEs	0.049	0.035	0.045	-8.55	21	1,293
EMEs	0.198	0.084	0.396	-8.50	25	1,354
Dlog real output						
All	0.040	0.040	0.041	-30.38	46	2,649
AEs	0.033	0.032	0.030	-19.25	21	1,293
EMEs	0.045	0.049	0.047	-13.55	25	1,356
Dlog money						
All	0.176	0.120	0.273	-16.37	46	2,563
AEs	0.091	0.085	0.060	-12.75	21	1,277
EMEs	0.261	0.175	0.362	-10.56	25	1,286
Dlog credit						
All	0.183	0.133	0.288	-18.69	46	2,563
AEs	0.104	0.102	0.071	-14.50	21	1,292
EMEs	0.263	0.182	0.387	-12.12	25	1,271

Table 1. **Descriptive statistics.** The table reports the mean, the median, the standard deviation (S.D.), the number of countries (N), the total number of observations (NxT) and the result of the Im-Pesaran-Shin (2003) panel unit root test for the full set of countries (All) as well as for the group of advanced economies (AEs) and emerging market economies (EMEs) separately. All unit root test statistics are significant at the 1% level.

median growth rates and a much lower standard deviation, both in advanced and emerging market economies. The average growth rate of real GDP across all countries over the post-war period was 4%. In the EMEs, it was 4.5%, more than 1 percentage point higher than in the AEs (3.3%).

The characteristics of the descriptive statistics of money and credit growth are similar to that of inflation. The mean growth rates are in both cases considerably larger than the median growth rates, indicating a number of large positive outliers. This feature is again driven by the EMEs, which have considerably higher means, medians and standard deviations for the two series. From a quantity theoretic perspective, the numbers in Table 1 roughly add up. For the full set of countries, the sum of the log change in the GDP deflator and of the log change in real GDP equals 16.5% while the average change in log broad money was 17.6%. This implies an average decline in the velocity of money of about 1.1% per year in our set of countries over the post-war period.

Credit has expanded more rapidly than money over the post-war period. For the full group of countries, the average and median growth of money was about 17.6% and 12% respectively, compared to 18.3% and 13.3% for credit. The difference was larger in AEs, where the mean and median growth rates of credit were respectively 1.3 percentage points and 1.7 percentage points higher than those of broad money. In EMEs, the difference in mean growth rates is only 0.2 percentage points per annum, while that in median growth rates is 0.7 percentage points.

The seemingly small differences in the yearly growth rates of money and credit have given rise to large cumulative differences in the evolution of the two variables over the post-war period (Figure 3, left-hand panel). Since 1950, the credit-to-GDP ratio in AEs has almost quadrupled, while the broad money-to-GDP ratio has less than doubled. Schularick and Taylor (2012) documented this leveraging trend for a smaller group of industrialised countries. Here we show that it holds more generally, for a larger group of AEs and also for EMEs. However, in EMEs, the divergence of credit and money growth has been somewhat less pronounced, with the credit-to-GDP ratio increasing more than fivefold since 1950 and the money-to-GDP ratio also more than quadrupling.[14]

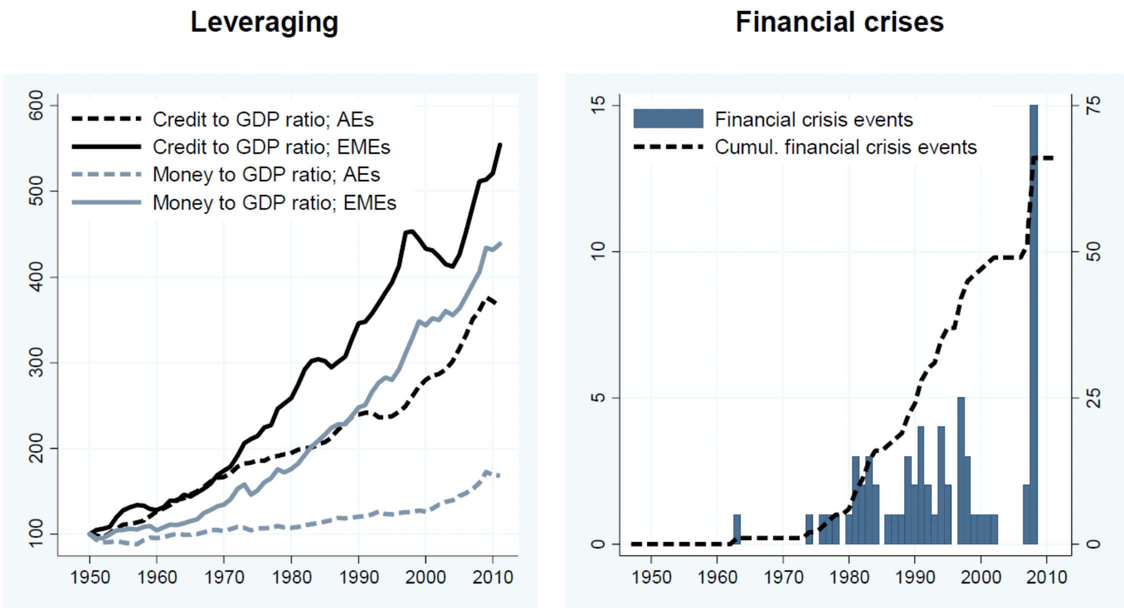

Figure 3. **Leveraging and financial crises.**

The rapid expansion in leverage was accompanied by growing financial stability risks. This is reflected in the growing incidence of financial crises since the 1980s, culminating in the Great Financial Crisis of 2008 affecting a large number of countries globally (Figure 3, right-hand panel). The total number of financial crises in our sample is 66.[15]

Finally, the unit root tests overwhelmingly reject the null hypothesis of a unit root for

[14] In emerging Asian economies, money-to-GDP ratios expanded faster and have often exceeded credit-to-GDP ratios since the late 1990s, reflecting the large stock of public securities held by the domestic banking sector as a consequence of the build-up of large FX reserve war chests in the wake of the Asian crisis (see Filardo et al. (2014)).

[15] Financial crises have occurred in 42 out of 46 countries in our sample. For more details, refer to Annex-Table A1.

all four variables (Table 1, column 5).[16] In the following empirical analysis we can therefore assume that we are dealing with stationary data series.

3 Empirical approach

In the empirical analysis, we have to take into account that the two monetary facts we revisit here are of different type and therefore require different econometric approaches. The money-inflation nexus represents a relationship between two continuous variables with a large number of cross-sectional and time series observations. The credit growth-financial crisis link, in contrast, represents a relationship between a continuous variable, credit growth, and a discrete event variable, the financial crisis dummy, with only a few positive (i.e. crisis) observations at the individual country level. Therefore, while the sufficient number of time-series observations allows potential cross-country heterogeneities in the money-inflation nexus to be taken into account, the limited number of crisis events at the individual country level precludes this for the exploration of the credit-crisis nexus.

The analysis of the money-inflation link follows McCallum and Nelson (2011) by allowing for a dynamic lead-lag relationship between money growth and inflation. To this end, we employ the pooled mean group (PMG) estimator developed by Pesaran et al. (1999). The PMG estimator is a maximum likelihood panel estimator constraining long-run coefficients to be the same and allowing short-run coefficients and error variances to differ across cross-sectional units.

The estimating equation is given by:

$$\Delta\pi_{i,t} = \alpha_i + \phi_i(\pi_{i,t-1} - \theta(m-y)_{i,t-1}) + \sum_{j=0}^{4} \delta_{i,j}\Delta(m-y)_{i,t-j} + \varepsilon_{i,t} \tag{1}$$

where π is the inflation rate measured as the log difference of the price level and m-y is excess money growth, measured as the log change in nominal money (m) less the log change in real GDP (y).

Equation (1) represents an error-correction model derived from a baseline autoregressive distributed lag (ARDL) model, linking inflation to current and lagged excess money growth and to its own lag.[17] Specifically, we include five lags of excess money growth and one lag

[16] The indications of the IMS test reported in Table 1 are robust to the use of alternative panel unit root tests, irrespective of whether they allow for individual unit root processes across cross-sections (like the IMS test) or whether they assume common unit root processes (such as e.g. the test proposed by Levin et al. (2002)). The results of these additional tests are available upon request.

[17] We model inflation as a function of money growth in excess of real GDP growth as this is consistent with the quantity equation and reduces the dimensionality of the model compared to the alternative approach of including nominal money growth and real GDP growth as separate regressors. In the robustness checks

of inflation in the ARDL model, thus allowing for long lead-lag relationships (up to five years) between the two variables. This way we also correct for potential endogeneity of the explanatory variable (Pesaran and Shin (1999)).

The second term in equation (1) represents the long-run relationship between inflation and excess money growth. It follows from log differencing the quantity equation $MV = PY$, which yields $\pi = m - y + v$. The quantity equation thus predicts a one-to-one relationship between inflation and the growth rate of money in excess of real GDP growth for given changes in velocity (v), i.e. the long-run coefficient θ is predicted to be equal to one. Trends and temporary fluctuations in velocity would be reflected in the country-fixed effects α_i and the country-specific error term $\varepsilon_{i,t}$.

The PMG estimator incorporates the assumption that the cross-sectional units (here countries) share the same long-run relationship, but differ in the short-term adjustments and dynamics around that long-run relationship. The long-term coefficient θ is thus restricted to be the same for all countries. ϕ_i is the country-specific error-correction coefficient. Panel estimates of this coefficient are obtained through the mean group procedure by averaging over the individual country estimates. The existence of a long-run relationship between inflation and money growth can be tested based on the t-statistic of the panel mean group estimate of ϕ.[18] The third term on the right-hand side of the equation captures the country-specific short-term dynamics. Specifically, $\delta_{i,j}$ are the country-specific coefficients for the current and lagged change in excess money growth. Panel estimates of these coefficients are also obtained through the mean group procedure.

The analysis of the link between credit growth and financial crisis is based on a probabilistic model linking the incidence of financial crises to lagged credit growth, following Schularick and Taylor (2012). Specifically, we estimate a panel model of the form:

$$I_{i,t} = \alpha + \sum_{j=1}^{5} \beta_j (c - \pi - y)_{i,t-j} + \eta_{i,t} \tag{2}$$

where $I_{i,t}$ is the financial crisis indicator taking the value one when a financial crisis occurred in country i in period t, and zero otherwise (as described in the data section), $c - \pi - y$ is real excess credit growth, i.e. the log change in nominal bank credit (c) less the log change in nominal GDP ($\pi + y$).[19] $\eta_{i,t}$ is the country-specific error term and α is

section we also report results for the specification using nominal money growth.

[18] Pesaran et al. (2001) show that the existence of long-run level relationships between two or more variables is not limited to the case of I(1) variables, but also applies to the case of I(0) variables. Specifically, they derive two sets of critical values of the F-and t-statistics of the error-correction term to test for the existence of a long-run relationship, one for the case of I(1) variables and one for the case of I(0) variables. According to the unit root tests reported in Table 1, the variables in our analysis are I(0).

[19] In the robustness checks section we also consider alternative credit-based crisis indicators, such as real

a regression constant. We include five lags of real excess credit growth taking into account that, as pointed out by Schularick and Taylor (2012), credit booms are phenomena that last several years.

We estimate equation (2) using a panel probit model, which performed somewhat better in terms of predictive ability than the alternative logit model (see the robustness checks section). The regression equation does not contain country-fixed effects since a number of countries did not experience a financial crisis over the post-war period. The count of these countries further increased when we re-estimated the model over sub-samples in a later stage of the analysis. If the models were estimated by standard fixed effects, the countries with no crisis events (i.e. with $I_{i,t} = 0$ over the entire estimation sample) would drop out of the regression. That would mean that only countries that experienced a crisis would be included in the analysis. In other words, we would incur a selection bias. In order to address this caveat, we measure real excess credit growth as a deviation from its respective country-specific mean and estimate the panel regression in pooled form.[20]

From estimating equation (2), we can assess the link between lagged excess credit growth and financial crisis events through the coefficients λ_j. Specifically, we can calculate the marginal effect of lagged credit growth on crisis probability (evaluated at the mean of the explanatory variable), which will tell us the impact of an increase in credit over GDP on the future probability of a financial crisis (in percentage points).

In order to assess the model's predictive ability, we calculate the AUROC, the area under the receiver operator curve (ROC). The ROC gives the combinations of true and false crisis calls for different call thresholds of the estimated models. More specifically, for the predicted crisis probability of the model we vary the threshold for which this probability would trigger a crisis call for all values between 0 and 1 and then obtain a curve of all true and false crisis calls for the different thresholds. The model's predictive ability can then be compared with an uninformed model that would always have the same number of true and false calls without the need to take a stand on the appropriate level of the threshold. Specifically, the AUROC of a given model can be tested against the null of 0.5, which would be the AUROC of the uninformed benchmark, based on the assumption of an asymptotic normal distribution.[21]

Equations (1) and (2) represent the baseline models that we use in the empirical analysis. There are, of course, alternative ways to specify the equations. There are, for instance,

credit growth and the credit gap.

[20] As explained in Schularick and Taylor (2012), we cannot estimate equation (2) with time-fixed effects as the model can then only be estimated using years with variation in the outcome variable, which dramatically reduces the number of observations.

[21] The AUROC has become widely used in the recent crisis prediction literature (e.g. Schularick and Taylor (2012), Drehmann and Juselius (2013)).

different possibilities to specify the dependent or the explanatory variable. And there is potential omitted variable bias as other factors might influence inflation and crisis probability besides money and credit growth. In Section 5 we assess the robustness of the baseline results to these caveats by performing a large number of robustness checks.

4 Post-war monetary facts: the evidence

4.1 Baseline facts

We start by exploring the two monetary facts, using all available observations over the post-war era. This gives us around 2,300 observations to estimate equations (1) and (2) respectively. The results reported in Table 2 confirm that there is a significant long-run link between money growth and inflation, and that credit growth has significant predictive power for financial crises.

Money-inflation nexus		**Credit-crisis nexus**	
Long run	0.961*** (0.012)	Overall ME	0.232*** (0.050)
Error correction	-0.495*** (0.042)	AUROC	0.698*** (0.037)
Short-term dynamics		ME of individual lags	
L0	-0.192*** (0.034)	L1	0.076*** (0.024)
L1	-0.121*** (0.030)	L2	0.065*** (0.022)
L2	-0.074*** (0.020)	L3	0.074*** (0.028)
L3	-0.053** (0.022)	L4	0.014 (0.033)
L4	-0.034** (0.015)	L5	0.004 (0.026)
RMSE	0.324	Pseudo R^2	0.060
Loglikelihood	-4610.4	Pseudolikelihood	-272.2
Observations	2,336	Observations	2,331
Countries	46	Countries	46

Table 2. **Full sample results.** The left-hand columns report the long-run coefficient and the error-correction coefficient from the Pooled Mean Group (PMG) estimation of equation (1). L0 to L4 refers to the coefficients of the current and lagged changes in excess money growth. RMSE is the root mean squared error. In the right-hand columns, the table reports marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel probit estimation of equation (2). Overall ME refers to the sum of the marginal effects of the five lags of credit growth. L1 to L5 are the marginal effects of the individual lags. Standard errors are in parentheses. Coefficient and ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

Money growth and inflation are closely linked in the long run with a long-run effect of money growth on inflation of 0.96 (Table 2, left-hand columns), which is very close to the one-to-one relationship predicted by the quantity theory. It is, however, still significantly smaller than one due to the very small standard error of the long-run coefficient. There is also strong and highly significant error correction of inflation to the long-run relationship with money growth (t-statistic = -11.80), suggesting that a long-run relationship between the two variables exists.[22]

The overall marginal effect of lagged credit growth (evaluated at its mean) on financial crisis probability is 0.23 and significant at the 1% level (Table 2, right-hand columns). This means that an increase in the credit-to-GDP ratio by 10% over five years increases the probability of a financial crisis by roughly 2 percentage points. This is a fairly high sensitivity given that the crisis frequency in our sample is around 3%. The AUROC statistic is 0.70 and significantly different from 0.5, suggesting that the model has fair predictive power.

4.2 Changes over time

To assess potential changes in the two monetary facts over time, we re-estimate equations (1) and (2) over sub-sample periods. Specifically, we consider potential structural changes to have occurred in the mid-1980s when financial liberalisation and disinflation took hold; and in the mid-1990s due to further disinflation and financial liberalisation in the late 1980s/early 1990s.

Table 3 reports the results for three sub-samples, 1950-1984, 1985-2011 and 1995-2011. The estimation results for the sub-sample 1950-1994 are reported for completeness and reference, if needed, in Annex-Table A2. The results suggest that the long-run link between money growth and inflation has weakened over time, while the link between credit growth and financial crisis has strengthened. Specifically, the long-run coefficient of money growth is estimated at 0.80 over the early sub-sample, not too far from but still significantly below the unit coefficient predicted by the quantity theory. It then drops to 0.21 in the sample period since 1985, and further halves to 0.11 when the sample begins in 1995 (Table 3, left-hand columns).

The finding that the long-run coefficient in the early sample period is smaller than over the full sample is due to the exclusion of a number of hyperinflation episodes in the late 1980s and early 1990s. If the model is estimated over the period 1950-1994 when these episodes are included, the long-run coefficient is estimated at 0.96 (see Annex-Table A2, left-hand columns).

Interestingly, the error-correction term remains highly significant throughout, also for the post-1994 sub-samples, when the long-run coefficient of money growth drops to very low levels. There is hence no error correction between inflation and money growth. Instead,

[22] The 1% critical value of the error-correction-based t-test for the existence of a level relationship model is -3.43 (Table CII in Pesaran et al. (2001)).

there is mean reversion of inflation to a country-fixed effect representing the country-specific explicit or implicit inflation target.

Money-inflation nexus				**Credit-crisis nexus**			
	1950-1984	**1985-2011**	**1995-2011**		**1950-1984**	**1985-2011**	**1995-2011**
Long run	0.797*** (0.035)	0.207*** (0.031)	0.115*** (0.018)	Overall ME	0.073 (0.055)	0.397*** (0.077)	0.415*** (0.092)
Error correction	-0.565*** (0.042)	-0.393*** (0.046)	-0.631*** (0.054)	AUROC	0.697*** (0.065)	0.729*** (0.038)	0.753*** (0.043)
Short-term dynamics				ME of individual lags			
L0	-0.159*** (0.040)	0.131*** (0.048)	0.021 (0.032)	L1	0.034 (0.023)	0.109*** (0.036)	0.167*** (0.046)
L1	-0.115*** (0.045)	0.036 (0.027)	-0.006 (0.022)	L2	0.034 (0.025)	0.114*** (0.037)	0.024 (0.045)
L2	-0.057 (0.043)	-0.004 (0.020)	-0.008 (0.026)	L3	0.024 (0.036)	0.116*** (0.042)	0.124*** (0.052)
L3	-0.013 (0.040)	-0.026 (0.018)	-0.025 (0.021)	L4	-0.049* (0.025)	0.070 (0.049)	0.116*** (0.047)
L4	-0.034 (0.033)	-0.003 (0.016)	-0.022 (0.021)	L5	-0.031 (0.027)	-0.012 (0.036)	-0.016 (0.037)
RMSE	0.346	0.221	0.112	Pseudo R^2	0.054	0.090	0.100
Log-Likelihood	-4470.0	-5802.6	-4500.6	Pseudolikelihood	-79.41	-180.6	-114.3
Observations	1,128	1,206	776	Observations	1,123	1,206	776
Countries	46	46	46	Countries	46	46	46

Table 3. **Changes over time.** The left-hand columns report the long-run coefficient and the error-correction coefficient from the Pooled Mean Group (PMG) estimation of equation (1). L0 to L4 refers to the coefficients of the current and lagged changes in excess money growth. RMSE is the root mean squared error. In the right-hand columns, the table reports marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel probit estimation of equation (2). Overall ME refers to the sum of the marginal effects of the five lags of credit growth. L1 to L5 are the marginal effects of the individual lags. Standard errors are in parentheses. Coefficient and ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

We also see considerable changes in the credit-crisis nexus (Table 3, right-hand columns). The marginal effect of lagged credit growth on crisis probability is almost twice as large in the more recent period beginning in 1985 compared to the full sample (0.40 vs 0.23). Compared to the pre-1985 sample the difference is even more striking. Over the early sample period 1950-1984, the overall marginal effect is estimated at 0.07 and is not statistically different from zero. In other words, the relationship between credit growth and financial crisis has emerged in a statistically significant way only over the past three decades. Previously it was absent, reflecting the absence of widespread financial fragility in the form of systemic crises. The aggregate marginal effect of lagged credit growth over the sample starting in 1995 is basically the same as that for the sample starting ten years earlier, so there do not seem to have been major further changes in the relationship since the mid-1990s.

We also see an improvement in the predictive ability of the panel probit models in the recent sample periods. The AUROC increases to 0.73 in the sample starting in the mid-80s, and then further to above 0.75 in the sample containing only observations since the mid-1990s. However, the standard error of the statistic is around 0.04, so that the difference from the full sample AUROC of 0.70 is not statistically significant.

The significance of the changes that we see over time is visualised in Figure 4, which shows the evolution of the long-run association between money growth and inflation and the marginal effect of credit growth on financial crisis probability over the three sub-samples. The chart shows how the money-inflation link weakens while the credit-crisis link strengthens, almost like a mirror image.

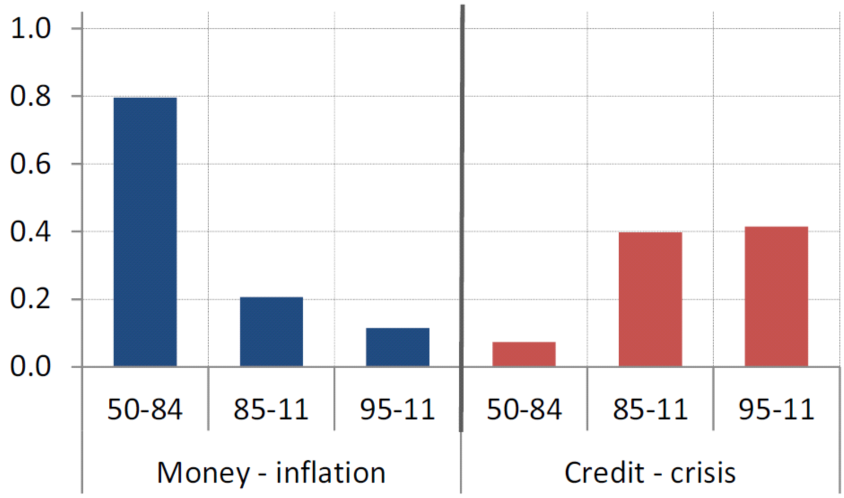

Figure 4. **Monetary facts over time.** Long-run impact of money growth on inflation and marginal effect of credit growth on crisis probability based on the estimates reported in Table 3.

4.3 Advanced vs emerging market economies

To assess whether the differences in the financial and monetary regime between advanced and emerging economies are reflected in the monetary facts, we split our sample of countries into a group of AEs, with 21 economies, and a group of EMEs, with 25 economies. We then re-estimate equations (1) and (2) for the two groups separately.

The full sample results, shown in Table 4, suggest that there are indeed significant differences between the two groups over the post-war period. In the AEs, the long-run coefficient of money growth over the full sample equals 0.63, while in the EMEs the quantity theoretic one-to-one relationship between money growth and inflation holds (Table 4, left-hand columns). For the relationship between credit growth and financial crises we find the reverse pattern (Table 4, right-hand columns). The marginal effect of credit growth on crisis probability is about twice as large in the AEs (0.38) as it is in the EMEs (0.21). Also the AUROC is slightly higher in the AEs (0.73), than in the EMEs (0.70).

Money-inflation nexus			**Credit-crisis nexus**		
	AEs	**EMEs**		**AEs**	**EMEs**
Long run	0.637*** (0.052)	0.974*** (0.012)	Overall ME	0.377*** (0.093)	0.206*** (0.060)
Error correction	-0.360*** (0.047)	-0.630*** (0.051)	AUROC	0.733*** (0.052)	0.697*** (0.047)
Short-term dynamics			ME of individual lags		
L0	-0.092*** (0.018)	-0.218*** (0.057)	L1	0.070 (0.058)	0.076*** (0.027)
L1	-0.123*** (0.018)	-0.069 (0.050)	L2	0.105** (0.049)	0.059** (0.025)
L2	-0.047*** (0.017)	-0.056* (0.032)	L3	0.205*** (0.057)	0.057* (0.032)
L3	0.002 (0.024)	-0.072** (0.036)	L4	-0.009 (0.045)	0.012 (0.038)
L4	-0.004 (0.016)	-0.043* (0.023)	L5	0.007 (0.037)	0.002 (0.031)
RMSE	0.064	0.457	Pseudo R^2	0.088	0.063
Log-Likelihood	-2939.0	-1684.0	Pseudolikelihood	-121.0	-146.5
Observations	1,181	1,155	Observations	1,192	1,939
Countries	21	25	Countries	21	25

Table 4. **Advanced vs emerging market economies.** The left-hand columns report the long-run coefficient and the error-correction coefficient from the Pooled Mean Group (PMG) estimation of equation (1). L0 to L4 refers to the coefficients of the current and lagged changes in excess money growth. RMSE is the root mean squared error. In the right-hand columns, the table reports marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel probit estimation of equation (2). Overall ME refers to the sum of the marginal effects of the five lags of credit growth. L1 to L5 are the marginal effects of the individual lags. Standard errors are in parentheses. Coefficient and ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

The estimation results for the two country groups for the sub-samples 1950-1984, 1985-2011 and 1995-2011 are shown in Tables 5 and 6, those for the sub-sample 1950-1994 are again reported in Annex-Table A2. The results suggest that the money-inflation link was already weakening in the advanced economies in the mid-1980s, with the long-run coefficient dropping to 0.15, and further to 0.11 in the post-1994 sample (Table 5, left-hand columns). In the EMEs, by contrast, the quantity theoretic one-to-one relationship between money growth and inflation holds also over the sample beginning in the mid-1980s (Table 5, right-hand columns). The link has only broken down since the mid-1990s, when the long-run coefficient dropped to 0.1. This finding is consistent with the notion that it is the monetary regime in particular that matters for the visibility of the money-inflation nexus. The relationship between money growth and inflation in the two groups of countries respectively weakened when widespread disinflation across countries set in, a development that took place in the mid-1980s in the AEs, and in the mid-1990s in the EMEs.

	Advanced economies			**Emerging economies**		
	1950-1984	**1985-2011**	**1995-2011**	**1950-1984**	**1985-2011**	**1995-2011**
Long run	0.623***	0.151***	0.117***	0.933***	0.994***	0.100***
	(0.056)	(0.032)	(0.019)	(0.039)	(0.014)	(0.044)
Error correction	-0.473***	-0.428***	- 0.665***	-0.686***	-0.556***	-0.598***
	(0.058)	(0.068)	(0.088)	(0.060)	(0.069)	(0.068)
RMSE	0.047	0.023	0.022	0.568	0.534	0.150
Log-likelihood	-1478.8	-1757.4	-1254.2	-770.4	-1169.3	-1004.1
Observations	614	567	357	514	639	419
Countries	21	21	21	21	25	25

Table 5. **The money-inflation link in AEs and EMEs: sub-sample analysis.** The table reports the long-run coefficient and the error-correction coefficient from the Pooled Mean Group (PMG) estimation of equation (1). RMSE is the root mean squared error. Robust standard errors are in parentheses. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

For the relationship between credit growth and financial crises we find bigger changes over time in the advanced economies than in the EMEs (Table 6). The marginal effect of credit growth is essentially zero in advanced economies in the pre-1985 period, and then rises to 0.56 and 0.63 in the post-84 and post-94 sub-samples respectively (Table 6, left-hand columns). In the EMEs, the marginal effect of credit growth on financial crisis probability is 0.09 for the pre-1985 period and it is significantly different from zero, albeit only at the 10% level (Table 6, right-hand columns). It then rises to 0.21 in the post-1984 sub-sample, and goes up further to around 0.3 in the sub-sample covering the period since the mid-1990s. The increase in predictive ability over time, measured through the AUROC, is considerable in both groups. In advanced economies, the AUROC increases from 0.72 in the early sample to 0.81 in the later sample periods. In the EMEs, it rises from 0.69 over the early sample to

0.75 and then on to 0.80 in the two later sub-sample periods respectively.

	Advanced economies			**Emerging economies**		
	1950-1984	**1985-2011**	**1995-2011**	**1950-1984**	**1985-2011**	**1995-2011**
Overall ME	0.017 (0.116)	0.562*** (0.156)	0.633*** (0.272)	0.092*** (0.077)	0.299*** (0.078)	0.313*** (0.093)
AUROC	0.719*** (0.141)	0.792*** (0.046)	0.796*** (0.058)	0.689*** (0.066)	0.747*** (0.054)	0.809*** (0.047)
Pseudo R^2	0.039	0.157	0.142	0.053	0.114	0.168
Pseudolikelihood	-23.25	-83.76	-53.4	-53.55	-87.71	-53.79
Observations	625	567	357	498	639	419
Countries	21	21	21	21	25	25

Table 6. **The credit-crisis link in AEs and EMEs: sub-sample analysis.** The table reports overall marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel probit estimation of equation (2). Standard errors are in parentheses. ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

Figure 5 summarises the main findings of this section's analysis in graphical form. It shows the long-run effect of money growth on inflation (left-hand panel) and the marginal effect of credit growth on crisis probability (right-hand panel) respectively for EMEs (red bars) and for AEs (blue bars) over the early sample 1950-1984 and the two more recent sub-samples 1985-2011 and 1995-2011. The chart shows how the money-inflation link has weakened over time, earlier in the advanced economies than in the EMEs; and how the credit-crisis nexus has become stronger at the same time, with a more significant change in the AEs compared to the EMEs.

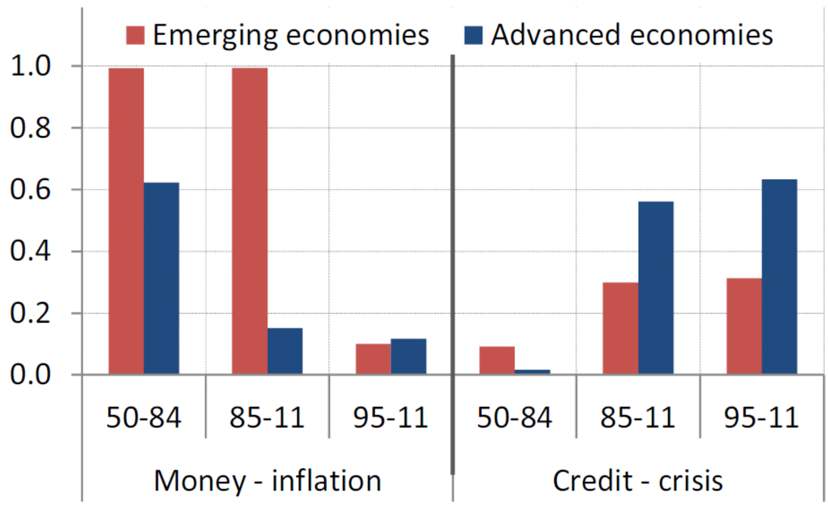

Figure 5. **Monetary facts in advanced and emerging economies over time.** Long-run impact of money growth on inflation and marginal effect of credit growth on crisis probability based on the estimates reported in Tables 5 and 6.

5 Robustness checks

In order to assess the robustness of the results of the previous section, we consider a number of alternatives to the baseline models.

5.1 Robustness checks for the money-inflation nexus

For the money-inflation link we consider the following robustness checks:

(i) We use the consumer price index (CPI) instead of the GDP deflator as the dependent variable. The GDP deflator is conceptually the right price measure to use in our analysis because it is the price index that relates to the measure of economic activity that is used (i.e. real GDP). Many studies have however used the CPI, which is commonly the prime gauge of price developments used by central banks.

(ii) We employ nominal money growth instead of excess money growth as the explanatory variable. This relaxes the assumption of a unitary income elasticity of money demand that is implicit in the use of the growth of money per unit of output.

(iii) We use M1 instead of broad money as the explanatory variable. M1 is a narrower monetary aggregate that is sometimes used instead of broad money in empirical studies (e.g. Teles et al. (2015)).

(iv) We adjust excess money growth for changes in velocity that can be linked to changes in the inflation rate. To this end, we first regress, for each country individually, the log change in velocity (calculated as the log change in nominal GDP less the log change in nominal broad money) on the change in inflation. Then we subtract the fitted values of this regression from excess money growth. The thus adjusted excess money growth series is then used as the explanatory variable in the PMG estimation. The adjustments are done separately for each of the sub-sample periods in order to take into account potential changes in the inflation elasticity of velocity over time.

This approach follows Teles and Uhlig (2013), except that we do the adjustment based on the change in inflation instead of the change in interest rates, given the lack of sufficiently long interest rate series for many EMEs. However, for the advanced economies the results from fitting velocity to the change in short-term interest rates instead of the change in inflation are very similar.[23]

(v) We assess the predictive power of money growth for inflation based on a mean group Granger causality test. Specifically, we regress inflation on its own lag and five lags of money growth country-by-county and calculate panel estimates of the coefficients using

[23] The results obtained based on this alternative approach to excess money growth adjustment are available upon request.

the mean group approach of Pesaran and Smith (1995), i.e. we average over the individual country estimates. The purpose of this exercise is to perform a cleaner test of the leading indicator property of money growth for inflation and its changes over time in order to see whether the same pattern holds as for the association between the two variables estimated through the PMG estimator.

(vi) We explore the relevance of omitted variable bias by including additional regressors in the PMG estimation. Specifically, we re-estimate the PMG model including also the output gap (estimated through a standard Hodrick-Prescott filter) and the log change in oil prices (West-Texas Intermediate).

For the sake of brevity, we report in Table 7 results only for the full set of countries and only for the full sample (1950-2011) and the two sub-samples 1985-2011 and 1995-2011.[24] The results for the estimations when we consider the groups of AEs and EMEs separately are reported in Annex-Table A3. In order to keep the size of the tables manageable, we report only a condensed regression output focusing on the key coefficients, i.e. the long-run coefficient and the error-correction term as well as the estimated coefficient for the additional variables when applicable. For ease of reference, we reproduce the baseline results at the bottom of the table.

Overall, the results suggest that the indications of the baseline analysis are robust. Specifically, the finding of a weakening of the long-run link between money growth and inflation over time generally obtains across all specifications, with long-run coefficients very similar to those obtained in the baseline model.

There are a number of further interesting results that stand out in Table 7. Narrow money growth displays a somewhat weaker association with inflation than broad money (rows under *3)* in Table 7). In particular in the advanced economies, the link between narrow money growth and inflation is weak over all sample periods (rows under *3)* in Annex-Table A3). In the EMEs, we also find that the quantity theoretic link was already weakening since the mid-1980s when it is tested based on narrow money growth. This suggests that broad money is a more adequate measure of monetary liquidity, consistent with the greater prominence of broad monetary aggregates in academic and central bank monetary analysis over the post-war period.

Adjusting excess money growth for inflation-driven changes in velocity raises the long-run coefficient somewhat in the post-1984 sample. It does however not restore the one-to-one relationship between money growth and inflation (rows under *4)* in Table 7).

The mean group Granger causality test shows that the predictive power of lagged money

[24] The results for the pre-1984 and pre-1994 sub-samples are available upon request.

growth for inflation has also faded, with the aggregate long-run multiplier of the lagged money growth terms falling from 0.47 in the full sample to 0.04 in the post-1994 sample period (rows under *5)* in Table 7). The estimate of the coefficient of lagged inflation also drops considerably over time, from 0.57 over the full sample to 0.29 in the post-94 sample. This confirms the finding established in the literature that inflation persistence is lower in low inflation environments (e.g. Benati (2008)).[25]

Finally, the regressions including either the output gap or the log change in the oil price show that the link between these two variables and inflation has also fallen over time (rows under *6)* and *7)* in Table 7). The long-run coefficient of the output gap has dropped from 0.93 in the full sample to 0.21 over the post-1994 sample, confirming the notion of a flattening of the Phillips curve in low inflation environments (e.g. Ball et al. (1988), Benati (2007)). Similarly, the long-run coefficient of the log change in oil prices has dropped from 0.08 to 0.02.

5.2 Robustness checks for the credit-crisis nexus

For the link between credit growth and financial crises, we consider the following robustness checks:

(i) We use the logit link function instead of the probit function. The baseline model in Schularick and Taylor (2012) is a fixed effects panel logit model.

(ii) We use real credit growth (i.e. the log change in nominal credit less the log change in the GDP deflator) instead of real excess credit growth (i.e. the log change in the credit-to-GDP ratio). Real credit growth is the indicator used by Schularick and Taylor (2012). In subsequent studies they have, however, also switched to real excess credit growth (Jordá et al. (2014)).

(iii) We employ one lag of the credit gap calculated as the difference between the credit-to-GDP ratio and its long-run trend estimated using a one-sided Hodrick-Prescott filter with a smoothing parameter of 1,600 as recommended by Drehmann et al. (2010).

(iv) We use a five-year moving average of excess credit growth as the dependent variable. This is the credit indicator used by Jordá et al (2014) as a practical alternative to the five-lags specification used in Schularick and Taylor (2012) for cases when more than one indicator is included in crisis prediction regressions.

(v) We extend the model to include additional regressors. Since we have only very few crisis observations, in particular for the sample period starting in 1995, we perform this robustness check as an extension of the previous one, i.e. using five-year moving averages

[25] The long-run multiplier is calculated as the sum of the coefficients of the money growth lags divided by one minus the coefficient of lagged inflation.

	1950-2011	1985-2011	1995-2011
1) PMG with CPI inflation			
Long run	0.981*** (0.014)	0.128*** (0.028)	0.050*** (0.020)
Error correction	-0.414*** (0.039)	-0.345*** (0.036)	-0.640*** (0.055)
2) PMG with nominal money growth			
Long run	0.971*** (0.015)	0.316*** (0.027)	0.141*** (0.021)
Error correction	-0.478*** (0.039)	-0.456*** (0.047)	-0.668*** (0.057)
3) PMG with narrow money growth			
Long run	0.785*** (0.035)	0.210*** (0.025)	0.128*** (0.021)
Error correction	-0.373*** (0.040)	-0.401*** (0.050)	-0.608*** (0.061)
4) PMG with velocity adjusted excess money growth			
Long run	0.969*** (0.012)	0.397*** (0.034)	0.128*** (0.018)
Error correction	-0.487*** (0.039)	-0.396*** (0.042)	-0.634*** (0.056)
5) MG Granger causality test			
Long run	0.468*** (0.009)	0.205*** (0.029)	0.040 (0.05)
Inflation persistence	0.570*** (0.068)	0.514*** (0.099)	0.285*** (0.060)
6) PMG adding output gap			
Long run money	0.955*** (0.013)	0.241*** (0.031)	0.089*** (0.020)
Long run output gap	0.929*** (0.102)	0.469*** (0.082)	0.212*** (0.047)
Error correction	-0.489*** (0.039)	-0.402*** (0.046)	-0.635*** (0.055)
7) PMG adding change in oil price			
Long run money	0.963*** (0.012)	0.284*** (0.040)	0.082*** (0.018)
Long run oil price	0.079*** (0.009)	0.038*** (0.008)	0.019*** (0.003)
Error correction	-0.486*** (0.039)	-0.330*** (0.036)	-0.525*** (0.052)
Memo item: Baseline PMG			
Long run	0.961*** (0.012)	0.207*** (0.031)	0.115*** (0.018)
Error correction	-0.495*** (0.042)	-0.393*** (0.046)	-0.631*** (0.054)

Table 7. **Robustness checks for the money-inflation link.** The table reports the long-run coefficient and the error-correction coefficient from the Pooled Mean Group (PMG) estimation of equation (1). In the case of the Granger causality test, the table reports the mean group estimate of the long-run multiplier of the lagged money growth terms, calculated as the sum of coefficients divided by one minus the coefficient of the lagged inflation term. Inflation persistence refers to the mean group estimate of the coefficient of the lagged inflation term. Robust standard errors are in parentheses. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

of the variables in order to avoid over-parameterisation of the model. Due to limited data availability over the post-war sample period in particular for many EMEs, we can consider only three standard additional crisis indicators: real GDP growth, GDP deflator inflation, and the change in the real exchange rate.

The exchange rate has been highlighted by Borio and Lowe (2002b) and Gourinchas and Obstfeld (2012) as an important indicator of financial distress, besides domestic credit expansion. Specifically, appreciation of the effective exchange rate has been found to indicate a growing risk of financial distress. Due to the unavailability of long series for effective exchange rates, we use instead the bilateral real exchange rate against the US dollar (for the United States we use an unweighted average of the bilateral exchange rate against the DM/euro and against the yen).[26]

The regressions again do not contain country-fixed effects since, as already discussed below, a number of countries did not experience a financial crisis over the post-war period or one of the sub-samples. Instead, we measure as before, the explanatory variables as a deviation from their country mean and then estimate the panel regression in pooled form. Also here we report, for brevity, in Table 8 only the results for the full set of countries and only for the sample periods 1950-2011, 1985-2011 and 1995-2011,[27] and focus on the key regression results, i.e. the overall marginal effect of credit on crisis probability, and on the AUROC. For the models considering additional regressors, we also report the marginal effect at the mean of the additional variables. The regression results for the separate groups of advanced economies and EMEs are reported in Annex-Table A4. Also here we show for convenience the baseline results at the bottom of the tables.

The results confirm the finding of the baseline analysis of an increase in marginal effects since the mid-1980s, and of enhanced predictive ability of the models over this period compared to the full sample going back to 1950. All the models yield similar marginal effects and AUROCs. Also the estimated increases over time in the marginal effect and the AUROC are very similar across models, in general confirming a doubling of the former and a slight improvement in the latter.

The results are also robust to the inclusion of additional indicators. Specifically, the marginal effects and the AUROCs of the moving average model are hardly affected by the inclusion of real GDP growth, inflation or the change in the real exchange rate. The marginal effect of credit on crisis probability is always significant at the 1% level and hardly changes

[26] In a recent paper, Hofmann et al. (2016) show that it is in fact the bilateral exchange rate against the US dollar rather than the effective exchange rate that matters for financial conditions in EMEs, reflecting liability dollarisation and dollar-based global bond investors.

[27] The results for the pre-1984 and pre-1994 sub-samples are available upon request.

	1950-2011	1985-2011	1995-2011
1) Logit			
Overall ME	0.204*** (0.047)	0.348*** (0.066)	0.349*** (0.078)
AUROC	0.700*** (0.037)	0.728*** (0.038)	0.746*** (0.044)
2) Probit with real credit growth			
Overall ME	0.192*** (0.047)	0.360*** (0.070)	0.387*** (0.084)
AUROC	0.673*** (0.037)	0.743*** (0.037)	0.759*** (0.042)
3) Probit with credit gap			
Overall ME	0.254*** (0.036)	0.332*** (0.053)	0.322*** (0.066)
AUROC	0.720*** (0.037)	0.730*** (0.037)	0.767*** (0.044)
4) Probit with 5-year MA of real excess credit growth			
ME	0.241*** (0.051)	0.412*** (0.080)	0.426*** (0.099)
AUROC	0.662*** (0.039)	0.707*** (0.039)	0.731*** (0.048)
5) Probit adding real GDP growth			
ME credit	0.249*** (0.051)	0.379*** (0.083)	0.365*** (0.105)
ME real GDP	-0.193 (0.146)	0.372 (0.294)	0.662* (0.371)
AUROC	0.670** (0.039)	0.715*** (0.039)	0.745*** (0.043)
6) Probit adding inflation			
ME credit	0.235*** (0.049)	0.426*** (0.079)	0.437*** (0.098)
ME inflation	0.035*** (0.010)	0.038** (0.015)	0.039 (0.027)
AUROC	0.690*** (0.037)	0.726*** (0.037)	0.746*** (0.043)
7) Probit adding change in real exchange rate			
ME credit	0.234*** (0.053)	0.393*** (0.081)	0.387*** (0.096)
ME exchange rate	-0.071 (0.071)	-0.110 (0.101)	-0.234** (0.107)
AUROC	0.675*** (0.038)	0.725*** (0.038)	0.746*** (0.041)
Memo item: Baseline probit			
Overall ME	0.232*** (0.050)	0.397*** (0.077)	0.415*** (0.092)
AUROC	0.698*** (0.037)	0.729*** (0.038)	0.753*** (0.043)

Table 8. **Robustness checks for the credit-crisis link.** The table reports marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel estimation of equation (2). Standard errors are in parentheses. ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

in magnitude across the various specifications.

For the additional indicators, we find that real GDP growth and GDP deflator inflation do not have a consistent significant impact on crisis probability (rows under *5)* and *6)* in Table 8). The change in the real exchange rate, by contrast, has a significantly negative effect on crisis probability over the post-1994 period (rows under *7)* in Table 8), suggesting that an exchange rate appreciation increases the probability of a crisis, consistent with the finding of Gourinchas and Obstfeld (2012). This result is mainly driven by the EMEs, where the exchange rate has a significantly negative effect that increases over time (rows under *7)* in Annex-Table A4). This finding is in turn consistent with the point made by Hofmann et al (2016) that an appreciation against the US dollar loosens financial conditions in EMEs, raising the risk of financial instability going forward.

6 Conclusions

This paper suggests that the long-run link between money growth and inflation, as well as the nexus between credit growth and future financial crises, have changed over time, but in reverse directions. While the former has weakened, the latter has strengthened. Moreover, we find that the money-inflation nexus has been stronger in EMEs than in advanced economies, while we find the opposite for the link between credit growth and financial crises. Against the background of the significant global disinflation and financial liberalisation trends since the mid-1980s and significantly lower inflation and more liberalised financial systems in advanced economies compared to the EMEs, these results indicate an antithetic relationship between the two monetary facts and the monetary and financial regime. While the money-inflation link is stronger in environments characterised by high inflation and low financial liberalisation, the reverse holds true for the credit-crisis link.

These results suggest that price stability and financial liberalisation could have implications for monetary analysis that go beyond the weakening of the link between money growth and inflation that was indicated by previous studies. Our analysis suggests that they also strengthen the link between credit expansion and financial crises. This could reflect the greater susceptibility of liberalised financial systems to generating credit boom-bust cycles that translate into greater financial fragility. Such instability may also be due in part to low inflation regimes underpinned by central bank credibility and global disinflationary forces. In such regimes, unsustainable economic and financial expansions appear to manifest themselves not primarily in inflationary pressures but instead in excessive credit growth and asset price booms that ultimately usher in financial crises (Borio (2014)).

The paper yields two further insights that are relevant for different strands of the literature. First, allowing for country-specific velocity trends and for lead-lag relationships between money growth and inflation, as we have done in this paper through the adoption of the Pooled Mean Group (PMG) estimator, does not overturn the result that the link between money growth and inflation has become weaker in countries with low inflation regimes. In other words, controlling for the two caveats that were emphasised by McCallum and Nelson (2011) does not restore the unitary relationship between money and inflation in a cross-country empirical set-up.

Second, the effect of credit expansion on financial crisis probability is probably considerably higher over the recent period than the estimates obtained in the literature based on long runs of historical data, in particular in advanced economies. Assessments of the benefits of integrating the credit-crisis link into monetary policy frameworks might therefore need to consider larger marginal effects of credit growth on crisis probability.[28]

[28] For instance, in the baseline panel fixed-effects logit model of Schularick and Taylor (2012) estimated on historical data for 14 advanced economies spanning the period 1870-2008, the marginal effect of credit growth on crisis probability is 0.3. This compares to a marginal effect of 0.65 in our baseline panel fixed effects probit model for advanced economies for the recent period 1995-2011.

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Appendix

	Money-inflation	Credit-crisis	Crises
United States	1950-2011	1950-2011	2
United Kingdom	1952-2011	1950-2011	4
Austria	1954-2011	1950-2011	1
Belgium	1954-2011	1954-2011	1
Denmark	1950-2011	1950-2011	2
France	1950-2011	1950-2011	1
Germany	1952-2011	1951-2011	1
Italy	1951-2011	1951-2011	2
Netherlands	1950-2011	1950-2011	1
Norway	1950-2011	1950-2011	1
Sweden	1950-2011	1950-2011	2
Switzerland	1950-2011	1950-2011	1
Canada	1950-2011	1950-2011	0
Japan	1954-2011	1954-2011	1
Finland	1950-2011	1950-2011	1
Greece	1954-2011	1954-2011	1
Ireland	1950-2011	1950-2011	1
Portugal	1954-2011	1954-2011	1
Spain	1953-2011	1950-2011	2
Australia	1950-2011	1950-2011	1
New Zealand	1950-2011	1950-2011	1
Turkey	1951-2011	1951-2011	2
South Africa	1966-2011	1966-2011	0
Argentina	1961-2011	1950-2011	4
Bolivia	1951-2011	1951-2011	2
Brazil	1950-2011	1950-2011	3
Chile	1962-2011	1962-2011	2
Colombia	1950-2011	1950-2011	2
Mexico	1950-2011	1950-2011	2
Paraguay	1951-2011	1953-2011	1
Peru	1950-2011	1950-2011	2
Uruguay	1950-2011	1950-2011	2
Venezuela	1950-2011	1950-2011	1
Israel	1961-2011	1961-2011	1
Hong Kong	1970-2011	1979-2011	0
India	1950-2011	1950-2011	1
Indonesia	1970-2011	1977-2011	1
South Korea	1954-2011	1962-2011	1
Malaysia	1966-2011	1966-2011	1
Philippines	1950-2011	1950-2011	2
Singapore	1964-2011	1964-2011	0
Thailand	1950-2011	1950-2011	2
Russia	1994-2011	1994-2011	2
China	1978-2011	1978-2011	1
Hungary	1991-2011	1991-2011	2
Poland	1991-2011	1991-2011	1

Table A1. **Country and time coverage of the analysis.**

	Money-inflation				Credit-crisis		
	All	AEs	EMEs		All	AEs	EMEs
Long run	0.958*** (0.017)	0.570*** (0.060)	0.976*** (0.018)	Overall ME	0.152*** (0.064)	0.309*** (0.089)	0.119 (0.084)
Error correction	-0.534*** (0.046)	-0.407*** (0.044)	-0.691*** (0.061)	AUROC	0.676*** (0.057)	0.772*** (0.074)	0.680*** (0.068)

Table A2. **Results for the sub-sample 1950-1994.** The left-hand columns report the long-run coefficient and the error-correction coefficient from the PMG estimation of equation (1). The right-hand columns report overall marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel probit estimation of equation (2). Standard errors are in parentheses. Coefficient and ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

	Advanced economies			Emerging economies		
	1950-2011	**1985-2011**	**1995-2011**	**1950-2011**	**1985-2011**	**1995-2011**
1) PMG with CPI inflation						
Long run	0.562***	0.088***	0.005	0.999***	1.052***	0.192***
	(0.062)	(0.033)	(0.020)	(0.014)	(0.014)	(0.024)
Error correction	-0.265***	-0.395***	-0.838***	-0.564***	-0.457***	-0.563***
	(0.019)	(0.041)	(0.072)	(0.055)	(0.071)	(0.079)
2) PMG with narrow money growth						
Long run	0.042	0.184***	0.101***	0.913***	0.604***	0.202***
	(0.084)	(0.026)	(0.024)	(0.028)	(0.068)	(0.039)
Error correction	-0.227***	-0.420***	-0.655***	-0.561***	-0.444***	-0.588***
	(0.035)	(0.078)	(0.110)	(0.050)	(0.063)	(0.067)
3) PMG with nominal money growth						
Long run	0.611***	0.249***	0.136	1.000***	1.002***	0.153***
	(0.038)	(0.032)	(0.026)	(0.014)	(0.017)	(0.034)
Error correction	-0.406***	-0.497***	-0.709***	-0.588***	-0.518***	-0.638***
	(0.052)	(0.064)	(0.088)	(0.054)	(0.067)	(0.076)
4) PMG with velocity adjusted excess money growth						
Long run	0.786***	0.192***	0.121***	0.979***	0.994***	0.165***
	(0.045)	(0.033)	(0.020)	(0.012)	(0.014)	(0.037)
Error correction	-0.370***	-0.443***	-0.722***	-0.599***	-0.528***	-0.571***
	(0.038)	(0.073)	(0.094)	(0.056)	(0.069)	(0.065)
5) MG Granger causality test						
Long run	0.510***	0.147***	0.060***	0.445***	0.248***	0.024*
	(0.005)	(0.006)	(0.003)	(0.022)	(0.079)	(0.014)
Error correction	0.667***	0.552***	0.297***	0.489***	0.481***	0.275***
	(0.050)	(0.066)	(0.091)	(0.117)	(0.176)	(0.082)
6) PMG adding output gap						
Long run money	0.651***	0.155***	0.090***	0.973***	0.994***	0.093
	(0.049)	(0.049)	(0.022)	(0.012)	(0.013)	(0.042)
Long run gap	1.406***	0.433***	0.176***	0.286**	0.259*	0.373***
	(0.152)	(0.100)	(0.049)	(0.143)	(0.167)	(0.118)
Error correction	-0.367***	-0.431***	-0.652***	-0.618***	-0.555***	-0.628***
	(0.037)	(0.064)	(0.090)	(0.054)	(0.072)	(0.069)
7) PMG adding change in oil price						
Long run money	0.512***	0.156***	0.079***	0.977***	0.999***	0.131***
	(0.052)	(0.041)	(0.020)	(0.012)	(0.014)	(0.046)
Long run oil price	0.105***	0.029***	0.018***	0.056***	0.085***	0.028***
	(0.011)	(0.007)	(0.003)	(0.014)	(0.019)	(0.012)
Error correction	-0.342***	-0.318***	-0.519***	-0.616***	-0.544***	-0.543***
	(0.044)	(0.035)	(0.083)	(0.048)	(0.065)	(0.068)
Memo item: Baseline PMG						
Long run	0.637***	0.151***	0.117***	0.974***	0.994***	0.100***
	(0.052)	(0.032)	(0.019)	(0.012)	(0.014)	(0.024)
Error correction	-0.360***	-0.428***	-0.665***	-0.630***	-0.556***	-0.598***
	(0.047)	(0.088)	(0.072)	(0.051)	(0.069)	(0.068)

Table A3. **Robustness checks for the money-inflation link in AEs and EMEs.** The table reports the long-run coefficient and the error-correction coefficient from the Pooled Mean Group (PMG) estimation of equation (1). In the case of the Granger causality test, the table reports the mean group estimate of the long-run multiplier of the lagged money growth terms, calculated as the sum of coefficients divided by one minus the coefficient of the lagged inflation term. Inflation persistence refers to the mean group estimate of the coefficient of the lagged inflation term. Robust standard errors are in parentheses. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

	Advanced economies			Emerging economies		
	1950-2011	**1985-2011**	**1995-2011**	**1950-2011**	**1985-2011**	**1995-2011**
1) Logit						
Overall ME	0.350***	0.485***	0.577**	0.184***	0.263***	0.262***
	(0.077)	(0.135)	(0.251)	(0.058)	(0.068)	(0.081)
AUROC	0.728***	0.792***	0.789***	0.696***	0.745***	0.802***
	(0.053)	(0.046)	(0.061)	(0.048)	(0.055)	(0.049)
2) Probit with real credit growth						
Overall ME	0.222***	0.465***	0.405*	0.182***	0.272***	0.284***
	(0.082)	(0.143)	(0.228)	(0.056)	(0.073)	(0.085)
AUROC	0.712***	0.796***	0.801***	0.673***	0.743***	0.822***
	(0.053)	(0.045)	(0.056)	(0.050)	(0.058)	(0.047)
3) Probit with credit gap						
Overall ME	0.216***	0.358***	0.391***	0.291***	0.311***	0.284***
	(0.045)	(0.082)	(0.118)	(0.057)	(0.068)	(0.079)
AUROC	0.749***	0.746***	0.724***	0.689***	0.715***	0.802***
	(0.061)	(0.053)	(0.081)	(0.048)	(0.053)	(0.043)
4) Probit with 5-year MA of real excess credit growth						
ME	0.418***	0.838***	0.975***	0.200***	0.284***	0.384***
	(0.107)	(0.204)	(0.388)	(0.066)	(0.097)	(0.101)
AUROC	0.686***	0.722***	0.675***	0.650***	0.669***	0.760***
	(0.059)	(0.052)	(0.083)	(0.055)	(0.068)	(0.061)
5) Probit adding real GDP growth						
ME credit	0.461***	0.755***	0.973***	0.212***	0.303***	0.265***
	(0.104)	(0.202)	(0.375)	(0.062)	(0.087)	(0.097)
ME real GDP	-0.468***	0.801	0.376	-0.084	0.217	0.661*
	(0.156)	(0.528)	(0.608)	(0.220)	(0.333)	(0.377)
AUROC	0.722***	0.743***	0.696***	0.649***	0.715***	0.812***
	(0.049)	(0.049)	(0.084)	(0.054)	(0.062)	(0.045)
6) Probit adding inflation						
ME credit	0.439***	0.913***	0.896***	0.199***	0.327***	0.351***
	(0.114)	(0.213)	(0.393)	(0.057)	(0.080)	(0.096)
ME inflation	0.024	0.269	-0.598	0.037***	0.033**	0.034
	(0.127)	(0.264)	(0.780)	(0.011)	(0.014)	(0.025)
AUROC	0.688***	0.727***	0.686***	0.704***	0.746***	0.799***
	(0.055)	(0.052)	(0.081)	(0.050)	(0.056)	(0.043)
7) Probit adding change in real exchange rate						
ME credit	0.426***	0.843***	0.958***	0.202***	0.298***	0.277***
	(0.108)	(0.201)	(0.386)	(0.064)	(0.084)	(0.090)
ME exchange rate	-0.209*	-0.154	0.015	-0.060	-0.107**	-0.209*
	(0.130)	(0.163)	(0.285)	(0.082)	(0.104)	(0.082)
AUROC	0.708***	0.728***	0.690***	0.658***	0.743***	0.828***
	(0.051)	(0.049)	(0.084)	(0.053)	(0.061)	(0.041)
Memo item: Baseline probit						
Overall ME	0.377***	0.562***	0.633**	0.206***	0.299***	0.313***
	(0.093)	(0.156)	(0.272)	(0.060)	(0.078)	(0.093)
AUROC	0.733***	0.792***	0.796***	0.697***	0.747***	0.809***
	(0.052)	(0.046)	(0.058)	(0.047)	(0.054)	(0.047)

Table A4. **Robustness checks for the credit-crisis link in AEs and EMEs.** The table reports marginal effects (MEs) evaluated at the mean and the area under the receiver operator curve (AUROC) calculated based on the panel estimation of equation (2). Standard errors are in parentheses. ME standard errors are robust. ***, ** and * denotes significance of a coefficient or test-statistic at the 1%, 5% and 10% level respectively.

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