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The Joint Dynamics of Money and Credit Multipliers Since the Gold Standard Era

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The Joint Dynamics of Money and Credit Multipliers Since the Gold Standard Era*

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Abstract

Since the XIX century, technological progress has allowed commercial banks to create ever greater amounts of broad money and credit starting from a unit of monetary base. Crucially, however, at the very low frequencies the *relative* amounts of the two aggregates created out of a unit of base money have remained unchanged over time in each of the 42 countries I analyze. This finding questions the widespread notion that, since WWII, credit has become disconnected from broad money, and suggests that, except for their greater productivity at creating broad money and credit out of base money, today's commercial banks are not fundamentally different from their XIX century's counterparts. The implication is that *only* the ascent of shadow banks has introduced a disconnect between broad money and credit.

Keywords: Money; credit; Lucas critique; financial crises.

*I wish to thank Peter Ireland for helpful discussions on money multipliers, suggestions, and comments; Juan-Pablo Nicolini for useful suggestions; Edward Nelson for comments, and Mark Watson for helpful suggestions on Müller and Watson's (2018) low-frequency regression methodology. Thanks to Alberto Musso and Athanassios Boudalis for kindly providing data for the Euro area and Greece, respectively. I am alone responsible for any mistake or imprecision.

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1 Introduction

The money multiplier, defined as the ratio between a specific monetary aggregate and the monetary base (e.g., M2 over M0), characterizes the amount of (broad) money that is created by the interaction between the financial system and the public out of a unit of base money provided by the monetary authority.[1] In a real sense it therefore characterizes a 'technology' that takes as 'input' a unit of base money, and produces as 'output' a specific amount of broad money.[2]

Money-multiplier analysis played a central role, first and foremost, in Milton Friedman and Anna J. Schwartz's (1963) *Monetary History of the United States*,[3] but over subsequent years and decades it was progressively abandoned, to the point that the money multiplier has almost faded from frontier research,[4] and the very concept has largely fallen into oblivion.[5]

In this paper I show that an analysis of the relationship between the multipliers of broad money and credit—with the credit multiplier being similarly defined as the ratio between a specific credit aggregate and the monetary base[6] (e.g., total loans to the private non-financial sector over M0)—sheds a new, and important light on the evolution of the global financial system since the Gold Standard era.

My main finding is that, since the XIX century, low-frequency fluctuations in the multipliers of broad money and credit have exhibited an extraordinarily strong correlation in each single one of the 42 countries I analyze. The long-horizon co-movement between the two multipliers has been so strong that, e.g., Wright's (2000) tests consistently detect *cointegration* between the two series, and the long-horizon correlation coefficient in the low-frequency regression of the credit multiplier on the money multiplier produced by Müller and Watson's (2018) methodology is most of the time close to one.

The simplest, and most powerful illustration of the strength and stability of the long-horizon relationship between the multipliers of broad money and credit is provided by the raw data showing their joint evolution, which are shown in Figures 1*a*-1*c*. The evidence speaks for itself: the low-frequency components of the two multipliers have been consistently moving in lockstep in all countries and nearly all periods,

[1] A small, but important qualification to this statement is that the monetary base can be defined even in the absence of a monetary authority (e.g., for the United States before the creation of the Federal Reserve system, see Friedman and Schwartz (1963)). This is why, throughout this paper, I also report a few results for samples during which a specific country did not have a central bank.

[2] See the discussion in (e.g.) Brunner and Meltzer (1990) and Modigliani and Papademos (1990).

[3] See, first and foremost, Chapter 7 ('The Great Contraction'). See also Phillip Cagan (1965).

[4] Among the very few recent theoretical analyses of models featuring a money multiplier, see Freeman and Kydland (2000) and Henriksen and Kydland (2010).

[5] E.g., whereas the 1990 *Handbook of Monetary Economics* featured extensive discussions of the money multiplier (see Brunner and Meltzer (1990) and Modigliani and Papademos (1990)), neither the subsequent 2011 edition, nor any edition of the *Handbook of Macroeconomics* even *mentioned* it.

[6] See e.g. Brunner and Meltzer (1990) and Modigliani and Papademos (1990).

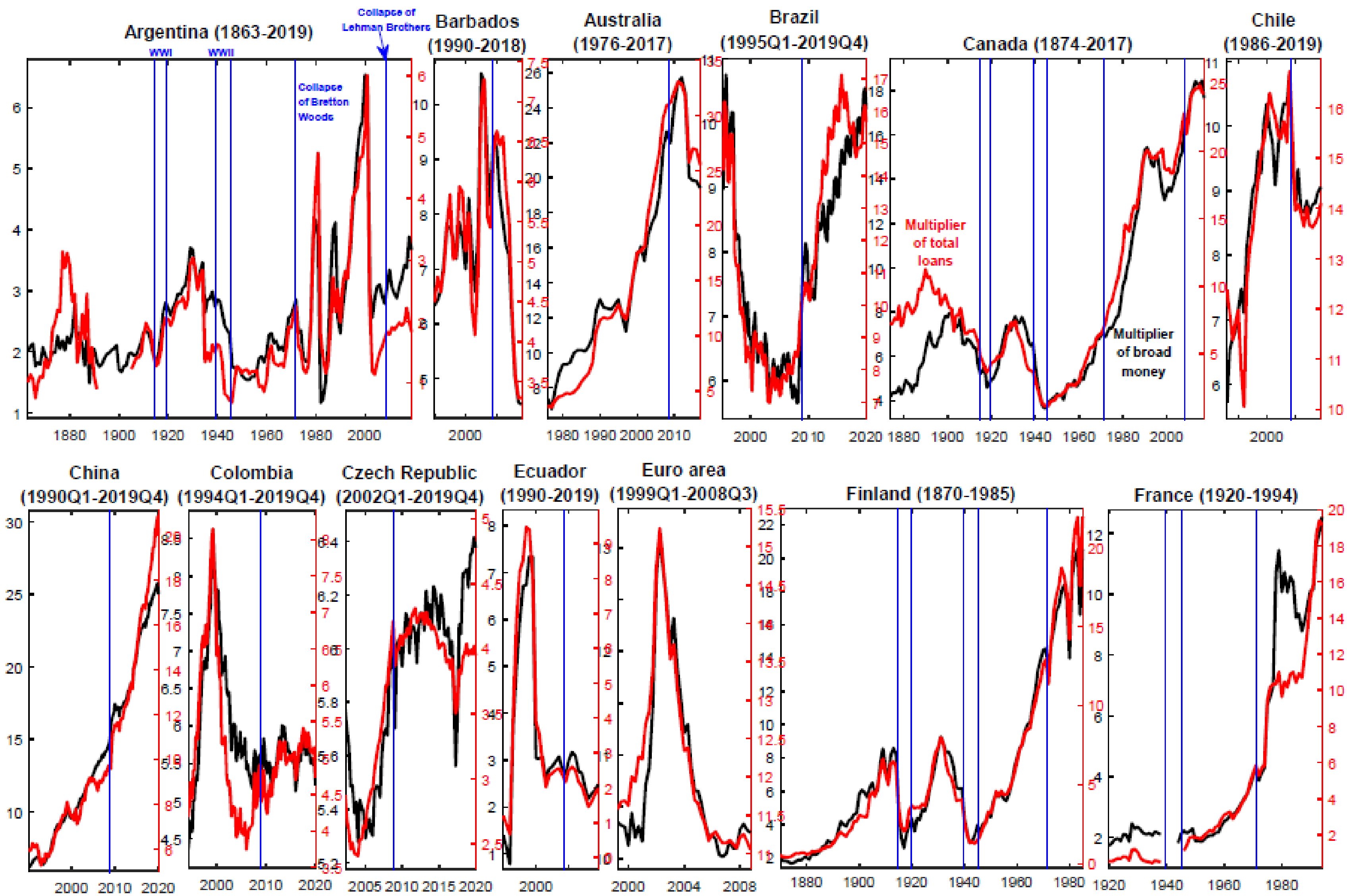

Figure 1a Multipliers of total loans and broad money for individual countries

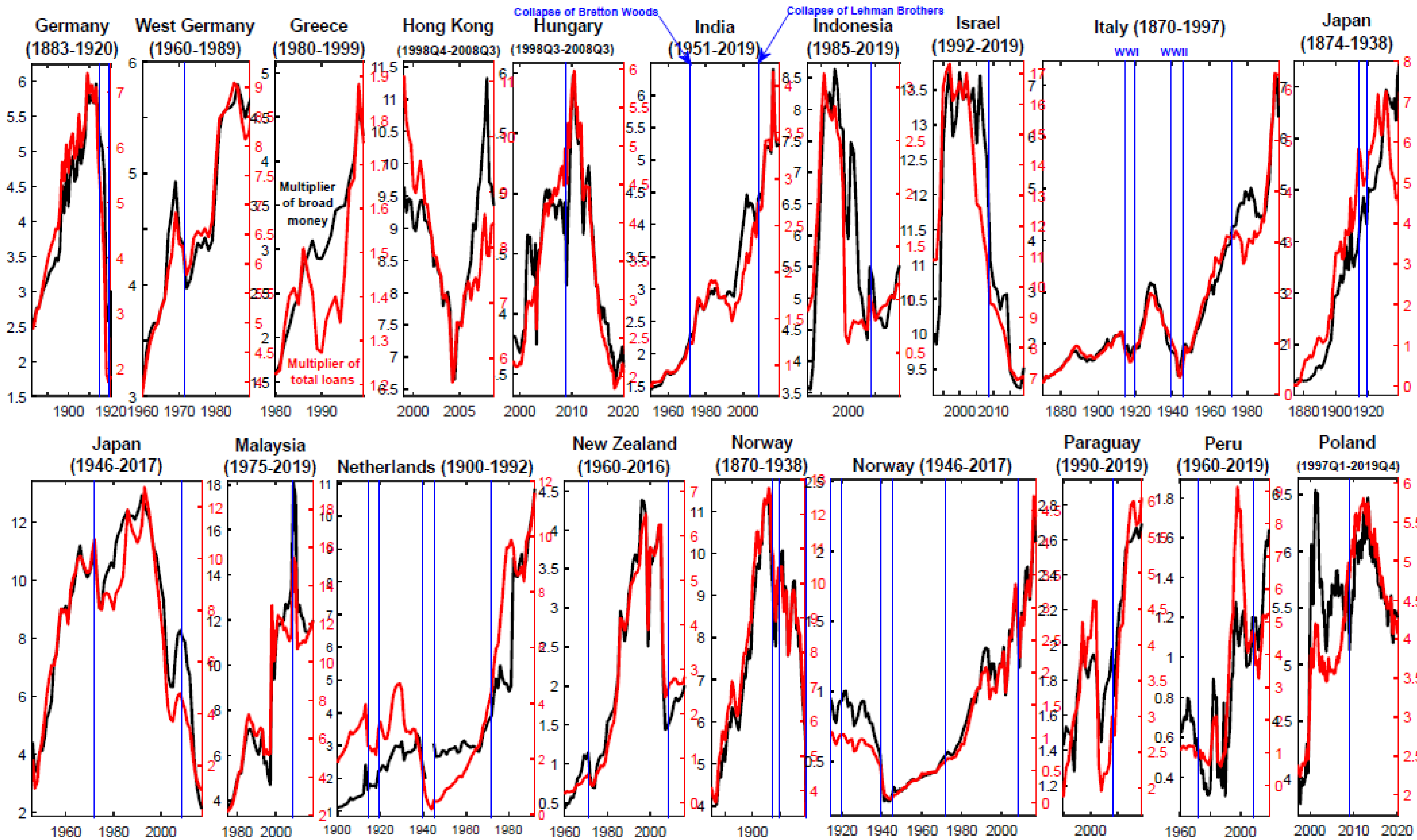

Figure 1b Multipliers of total loans and broad money for individual countries

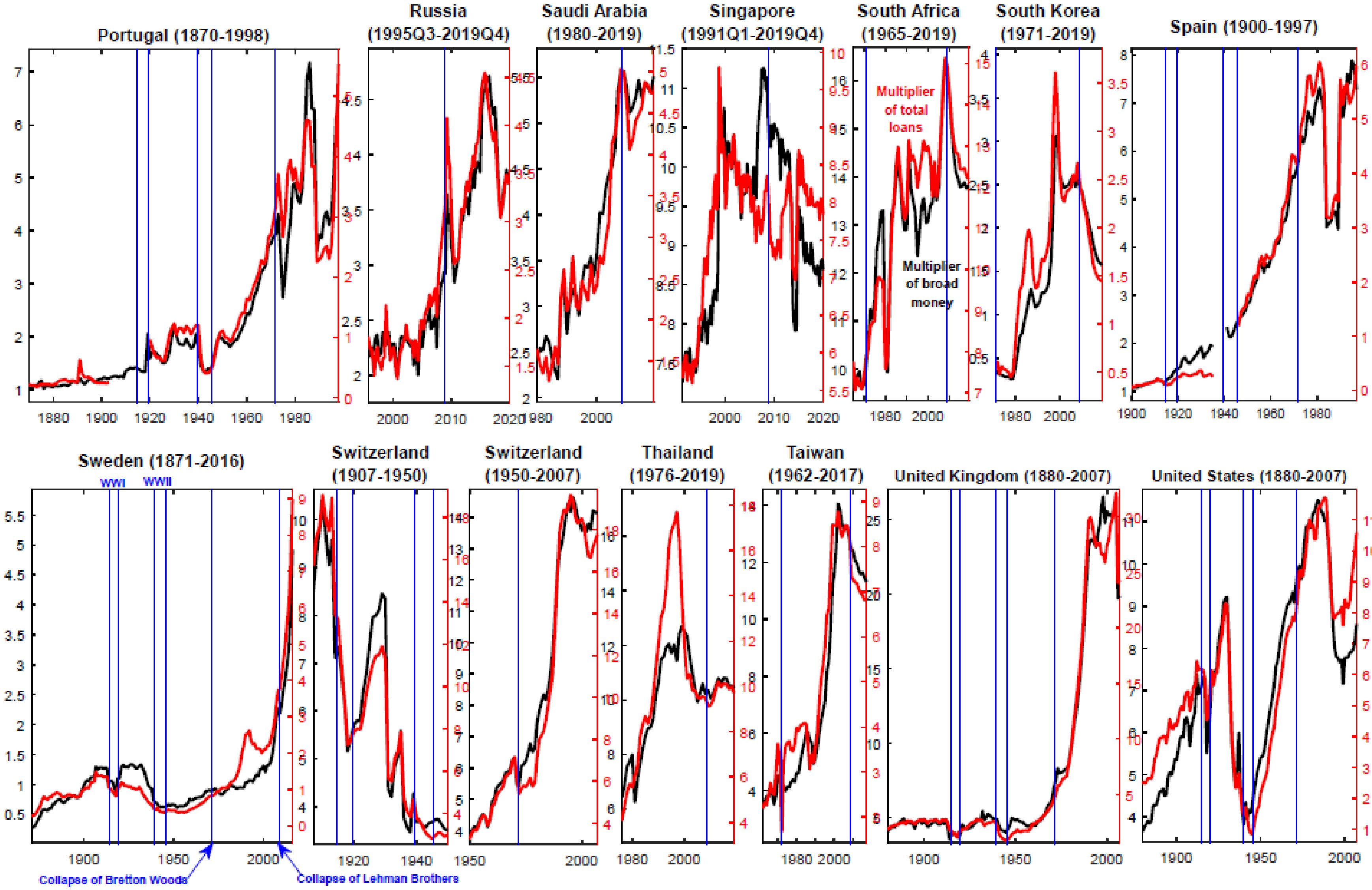

Figure 1c Multipliers of total loans and broad money for individual countries

to the point that in several cases they have been nearly indistinguishable.[7] This is the case, e.g., for Finland, Italy, Portugal, Spain, and the U.K. Statistical analysis, based on either country-and-year fixed-effects regressions, low-frequency regressions, or cointegration tests, will simply confirm what the visual evidence so starkly suggests.

This finding has a simple, and straightforward interpretation. Since (at least) the XIX century, technological progress has allowed commercial banks to create greater and greater amounts of both broad money and credit starting from a unit of base money. The evidence in Figures 1*a*-1*c* shows that this process has consistently taken place in a very specific and peculiar way: at long horizons, the *relative* amounts of the two aggregates created starting from a unit of base money have remained *unchanged* over time. A crucial point to stress is that the period since WWII has not exhibited *any* difference compared to the pre-WWII period. This also holds for the most recent years, encompassing the financial crisis and the Great Recession.

By defining the logarithms of the multipliers of broad money and credit as μ_M and μ_C, this evidence (as well as the econometric evidence in Section 3) suggests therefore that, to a first approximation, at long horizons $\mu_M = \alpha\gamma_t$ and $\mu_C = \beta\gamma_t$, with α and β being *constant* positive scalars, and γ_t capturing technological progress in the banking sector. This implies that, at long horizons, the creation of broad money and credit starting from a unit of base money has consistently proceeded in lockstep: in each country i, and nearly all periods, an increase in log broad money by one unit has been accompanied, in the long run, by an increase in log credit by β_i/α_i units, with α_i and β_i having remained constant over the entire sample period.

The implication is that, except for commercial banks' greater and greater productivity, the process that creates broad money and credit starting from a unit of base money *has not fundamentally changed since the XIX century.* This finding questions the widespread notion that, since WWII, and especially since the 1970s, credit has become progressively disconnected from broad money:[8] my evidence rather suggests that the only material change has been technological progress in the banking sector, which has affected broad money and credit in exactly the same way.

Based on the previous discussion, it is immediately apparent that the evidence that is typically used in support of the notion of a progressive disconnect between broad money and credit since WWII—i.e., the increase in the ratio between credit and money since 1945—is in fact uninformative. The reason is straightforward. Assume, just for the sake of the argument, that γ_t is a random-walk with drift,[9] so that the multipliers of broad money and credit are cointegrated. Unless $\alpha = \beta$, the logarithm of the ratio between credit and money, $\mu_C - \mu_M = (\beta - \alpha)\gamma_t$, features a unit root,

[7] Quite obviously, once appropriately rescaled. In Figures 1*a*-1*c*, the scales of the left- and right-hand side vertical axes are different, which amounts to linearly transforming the two series so that they have the same scale.

[8] See e.g. Schularick and Taylor (2012) and Jordà, Schularick, and Taylor (2017).

[9] The argument also holds if γ_t features a deterministic time trend.

thus spuriously pointing, under the standard interpretation of this metric, towards a disconnect between the two aggregates. In fact evidence clearly suggests that $\beta > \alpha$, so that increases in γ_t and/or the monetary base induce larger increases in credit than in broad money. In turn, this leads to an increase in the ratio between credit and broad money, thus giving the illusion, under the standard interpretation of this metric, of a progressive disconnect between the two aggregates. The logical implication is that the progressive increase in the ratio between credit and broad money since WWII[10] bears *no implication* for the issue of whether the two aggregates have, or have not become disconnected.

All of my evidence pertains to the 'traditional' (i.e., non 'market-based', or non 'shadow') banking sector. This is because, with the notable exception of the U.S., until quite recently data on shadow banks were not being collected systematically, thus making it impossible to perform a comparable analysis for overall credit in the economy (i.e., also including the loans created by shadow banks). Over the last four decades, however, the shadow banking system has experienced an extraordinary expansion.[11] For the present purposes, shadow banks feature a crucial difference with respect to the financial institutions that belong to the traditional banking sector. Since they finance the loans they create not by taking deposits from the public, but rather by borrowing on capital markets, such loans feature no corresponding monetary liability that could be counted as part of broad money. As a result, every time shadow banks generate a loan, they automatically introduce a 'wedge' (i.e., a disconnect) between broad money and credit, because they do not create any corresponding monetary liability.

My evidence therefore suggests that the *only* reason why, today, *overall* credit in the economy (i.e., credit generated by both shadow banks and traditional financial institutions) is partially disconnected from broad money is because of the dramatic expansion of shadow banking over the last four decades. If it had only been for the traditional financial institutions, there would be no disconnect between overall credit and broad money, and central banks, by steering broad money, would be able to indirectly reign in credit.

The paper is organized as follows. The next section provides a brief description of the data, which are discussed in detail in online Appendix A. In Section 3 I discuss evidence on the joint evolution of the multipliers of broad money and credit since the mid-XIX century from country-and-year fixed-effect regressions, Müller and Watson's (2018) low-frequency regression methodology, and Wright's (2000) cointegration tests. As I will discuss more extensively in Section 3.3 and 3.4, respectively, a key feature of both approaches is their flexibility. Müller and Watson's methodology has been designed to work well with series characterized by a wide array of low-frequency behaviour, from I(0) processes to (near) unit roots, to cointegrated processes. By the

[10] Documented e.g. by Jordà, Schularick, and Taylor (2017).

[11] See e.g. Adrian and Shin (2008, 2009, 2010, 2011). In Section 4.2 I discuss some evidence on this for the U.S.

same token, Wright's test has been designed to work equally well with either exact or near unit roots. Section 4 discusses several implications if these findings, in particular for macroeconomic analysis, macroeconomic policy, and for interpreting some of the previous evidence in the literature. Section 5 concludes, and outlines directions for further research.

2 The Data

Throughout the paper I report results based on two alternative datasets, a 'narrow' one that comprises 15 countries from Jordà, Schularick, and Taylor's (JST) dataset,[12] and a 'broader' one that also features 26 additional countries. All of the data and their sources are described in detail in online Appendix A. In this section I provide a brief overview of the main features of the two datasets, and of the data sources.

For the 15 countries in the narrow dataset, data on broad money, total loans (i.e., credit), nominal GDP, and the price level are all from JST's dataset. As for the monetary base, the narrow monetary aggregate featured in JST's dataset (i.e., the series labelled as 'narrowm') is equal to the monetary base only for Norway, Sweden, and the U.S., whereas it is equal to M1 for all other countries except the U.K., for which it is equal to 'Coins and notes in circulation'.[13] For Norway, Sweden, and the U.S. I have therefore taken the monetary base from JST's dataset, whereas for the remaining 12 countries I have taken it from national central banks' websites or statistical publications (for details, see online Appendix A). Finally, I eschew two countries in JST's database (Belgium and Denmark) because for either of them I was not able to find long-run series for the monetary base.[14]

As for the remaining 26 countries, data on total loans are in most cases from the *BIS*,[15] and they near-uniformly only cover the post-WWII period. Only for Argentina, I was able to extend the loans series back to 1863 based on the data featured in Ferreres (2005).[16] Data on broad money, nominal GDP, the price level, and the monetary base are all from either national central banks' websites and statistical publications, or national statistical agencies' websites.

[12] The dataset is available from the internet at: http://www.macrohistory.net/data/. The countries are Australia, Canada, Switzerland, Germany, Spain, Finland, France, the U.K., Italy, Japan, the Netherlands, Norway, Portugal, Sweden, and the U.S.

[13] See the description of JST's data in Jordà *et al.* (2021).

[14] For Denmark a series for the monetary base cannot be computed due to the lack of data on commercial banks' *reserves* (I wish to thank Kim Abildgren, of the Danish central bank, for confirming this to me). Interestingly, in line with the evidence in Figure 1, the ratios between either broad money or credit and *currency* (i.e., the second component of the monetary base beyond reserves) have exhibited an extraordinarily close correlation over the entire sample since 1875 (see Figure A.1 in the Online Appendix.

[15] See at: https://www.bis.org/statistics/totcredit.htm

[16] Over the period of overlapping (1941-2004) the series from the *BIS* dataset is near-identical to that from Ferreres (2005).

An important point to stress from the outset is that *all* of the credit (i.e., loans) data I work with throughout the paper *uniquely* cover the 'traditional' (i.e., 'non market-based', or 'non-shadow') banking sector. Shadow-banking, however, has exhibited an extraordinary growth over the last four decades. As already mentioned, it is this paper's main contention that only the ascent of shadow-banking has introduced a 'wedge' (i.e., a disconnect) between broad money and credit: as far as the 'traditional' banking system is concerned, my evidence suggests that, apart from their greater ability to transform base money into broad money and credit, today's banks are not fundamentally different from their XIX century's counterparts.

As for the sample periods, for each country I consider the longest available sample, with the single exception that for the U.S., the U.K., Switzerland and the Euro area, whose monetary policies following the financial crisis have led to dramatic increases in the monetary base, I end the sample periods in 2007. The reason for this is that including the subsequent period would distort the inference, since the dramatic increase in the monetary base mechanically caused a simultaneous collapse in the two multipliers, thus artificially 'blowing up' the strength of their correlation. On the other hand, for Japan I do not exclude the period of quantitative easing (QE) that started in early 2001, since the expansion in the monetary base was significantly more gradual. Finally, since within the European Monetary Union (EMU) the monetary base for individual countries is not defined, for all of these countries I end the samples at the latest in 1998.

Whereas for the countries in JST's dataset the sample periods typically start in 1870,[17] and for Argentina it starts in 1863, for several countries in the larger dataset the samples are quite short. This is the case, e.g., for Brazil, Colombia, and Russia, for all of which the sample starts in the mid-1990s, whereas for China it starts in 1990. In what follows, *all econometric work* for the larger dataset will be based on countries whose samples start at least in 1995, whereas I use the countries with shorter samples only for 'plotting' purposes, i.e. to visually illustrate the joint dynamics of the two multipliers over the most recent years. Although the samples starting after 1995 are in fact quite short, for this paper's purposes the evidence they provide is invaluable, because they clearly show that, contrary to the conventional-wisdom notion of a disconnect between broad money and credit since WWII, even in recent years long-horizon fluctuations in the two aggregates' multipliers have proceeded in lockstep.

For West Germany I restrict the sample to 1960-1989. The reason is discussed in detail in Online Appendix A.12.2 of Benati, Lucas, Nicolini, and Weber (2018), and it is briefly summarized in the online data Appendix A to the present work.[18] As for

[17] In a few cases data for the monetary base start much later (e.g., for Australia in 1976), thus compelling me to use shorter samples.

[18] In short, the data before 1960 did not include West Berlin and the Saarland (which in 1960 jointly accounted for about 6 per cent of overall GDP), whereas German reunification introduced discontinuities in the series.

Switzerland, since, as discussed in the online data Appendix A, the two series for the monetary base for the periods 1907-1950 and 1950-2006 cannot be linked, I consider the two periods separately.

Finally, throughout the entire paper I uniquely present results for either the *levels* or the *first differences* of either the multipliers of broad money and credit, or the ratios between the two aggregates and nominal GDP. The corresponding results for either the *logarithms* or the *log-differences* of either object, which are not reported for reasons of space, are qualitatively the same, and quantitatively very close. This entire, alternative set of results is available upon request.

I now turn to the evidence.

3 The Joint Evolution of Money and Credit Multipliers Since the XIX Century

In building up my argument that, since the second half of the XIX century, low-frequency fluctuations in the multipliers of broad money and credit have exhibited an extraordinarily strong correlation, I start from the simplest kind of evidence, i.e. the raw data, and I then move to progressively more sophisticated methods: country-and-year fixed-effects regressions; low-frequency regressions; and cointegration tests. In doing so I am motivated by the conviction, forcefully articulated (e.g.) by Summers (1991), that the most convincing type of empirical evidence is the simplest.

3.1 A look at the raw data

Figures 1*a*-1*c* show the joint evolution of the two multipliers for each of the 42 countries in the dataset. The evidence in the figures points towards a remarkably strong, and almost uniformly stable correlation between long-horizon fluctuations in the two series since the second half of the XIX century, thus illustrating how the 'production' of broad money and credit starting from a given 'input' of base money has consistently proceeded essentially in lockstep, especially at the very low frequencies.

An important point to stress is that such a strong and stable relationship between the two multipliers crucially hinges on working with the *overall* amount of credit granted by the traditional banking sector to the private non-financial sector. In contrast, splitting overall credit into (e.g.) either (*i*) a real estate and a non real estate segment, or (*ii*) segments pertaining to households and non-financial corporations, produces corresponding credit multipliers that, in general, exhibit a much weaker correlation with the multiplier of broad money than in Figures 1*a*-1*c*.[19] The implication is that the traditional banking sector's progressive shift towards real estate loans documented, e.g., by Jordà *et al.* (2016, Figure 4) has taken place, since WWII, in a very

[19] This evidence is available upon request. I do not report it simply because it is not especially interesting.

specific way. Based on the notation used in the Introduction, starting from a long-run equilibrium any increase in the amount of loans to real estate had to be consistently matched by a corresponding decrease in credit to the non real estate sector, so that the resulting log-multiplier for *overall* credit remained equal to $\mu_C = \beta\gamma_t$. To put it differently, *ceteris paribus* any shift in lending towards the real estate sector had to be accompanied by a corresponding contraction in lending to the non-real estate sector. The same logic holds for alternative ways of splitting overall credit, e.g. into portions pertaining to households and non-financial corporations.

It is worth briefly mentioning the very few cases in which the multipliers either temporarily diverge, or exhibit a somewhat weak correlation. Two interesting examples of temporary divergence are France and Italy during the Great Inflation episode. Whereas, for either country, the multipliers have evolved in lockstep over the rest of the samples, in both cases the Great Inflation has been characterized by a temporary, sizeable increase in the money multiplier compared to the credit multiplier. Although I have no explanation for such temporary divergence, the similarity between the two episodes, for two countries which had similar overall macroeconomic experiences during those years, naturally suggests that they might have been driven by the same mechanism. By the same token, an interesting feature common to both Argentina and Canada is that the two multipliers' fluctuations have been more strongly correlated *after* WWI than before, with the correlation having been remarkably strong in the latter period: this is exactly the *opposite* of the conventional-wisdom notion of a disconnect between broad money and credit since WWII. For Thailand the credit multiplier exhibits a large transitory increase, compared to the money multiplier, around the time of the 1997 Asian crisis. The same holds for Sweden around the time of the financial crisis of the early 1990s, and for the U.S. during the years leading up to the 2008-2009 financial crisis. By the same token, Singapore exhibits a deviation between the two multipliers around the time of the recent financial crisis.[20] Greece and Poland exhibit transitory deviations between the multipliers between the mid-1980s and the mid-1990s, and during the years leading up to the financial crisis, respectively. Finally, the Netherlands is the only country for which divergences between the two multipliers appear to have been quite frequent and long-lasting. From a close analysis of the evidence in Figure 1*b* it is however quite apparent how this is partly an illusion originating from the sizeable fall in the loans multiplier, compared to the money multiplier, during WWII. In fact, analyzing separately the two sub-samples before and after WWII, the correlation between the multipliers appears quite strong in both of them.

[20]The relationship appears to have reasserted itself since the mid-2010s, but with a shift in the intercept.

3.1.1 Summing up

The evidence in Figures 1a-1c shows that, since the second half of the XIX century, low-frequency fluctuations in the multipliers of broad money and credit have consistently exhibited a very strong correlation within each of the 42 countries I analyze. It is important to stress that

(i) by no means the correlation appears to have weakened after WWII: rather, in several cases (e.g., Argentina and Canada) it appears to have become *stronger*.

(ii) This stylized fact has consistently held both for the advanced countries in the narrow dataset, and for comparatively less developed countries such as Argentina, Brazil, Colombia, India, Malaysia, Russia, and China.[21]

(iii) By the same token, this has equally held for low-inflation countries such as Germany, Japan, and Switzerland; for countries that experienced sizeable inflation fluctuations during the Great Inflation episode, such as (e.g.) Italy, France and the the U.K.; and for very high-inflation countries such as Argentina, Brazil, and Israel.

The fact that the long-horizon relationship between the two multipliers has remained essentially unchanged during a period encompassing dramatic changes in the monetary regime—from the Gold Standard, to the introduction of unconventional monetary policies—naturally suggests that it is structural in the sense of the Lucas (1976) critique, thus reflecting fundamental structural features of how financial capitalism works. Further, the fact that the relationship has remained unchanged in the face of the shocks associated with the two World Wars, the Great Depression, and the 2008-2009 financial crisis shows that this is one of the most robust stylized facts in empirical macroeconomics. In fact, what could be deeper, more structural, and more robust than a relationship that has remained unaffected by *anything* that has happened since the second half of the XIX century?

I now turn to the evidence from country-and-year fixed-effects regressions.

3.2 Evidence from country-and-year fixed-effects regressions

The first two panels of Figure 2 show, based on either of the two datasets, the estimated year effects from country-and-year fixed-effects regressions for either of the two multipliers,[22] i.e.

$$\mu_{x,t}^{i} = a_i + b_t + e_{it} \tag{1}$$

where $\mu_{x,t}^{i}$ is the multiplier of either broad ($x = M$) money or credit ($x = C$) for country i; a_i and b_t are a country-specific and a year-specific fixed-effect, respectively; and e_{it} is an error term. I estimate (1) *via* OLS.

[21] In the early part of the sample (i.e., the 1990s) China was significantly less developed.than in the 2010s. In spite of this, the relationship between the two multipliers appears to have been the same throughout the entire sample.

[22] For all countries for which the data are available at the quarterly frequency, I convert them to the annual frequency by taking annual averages.

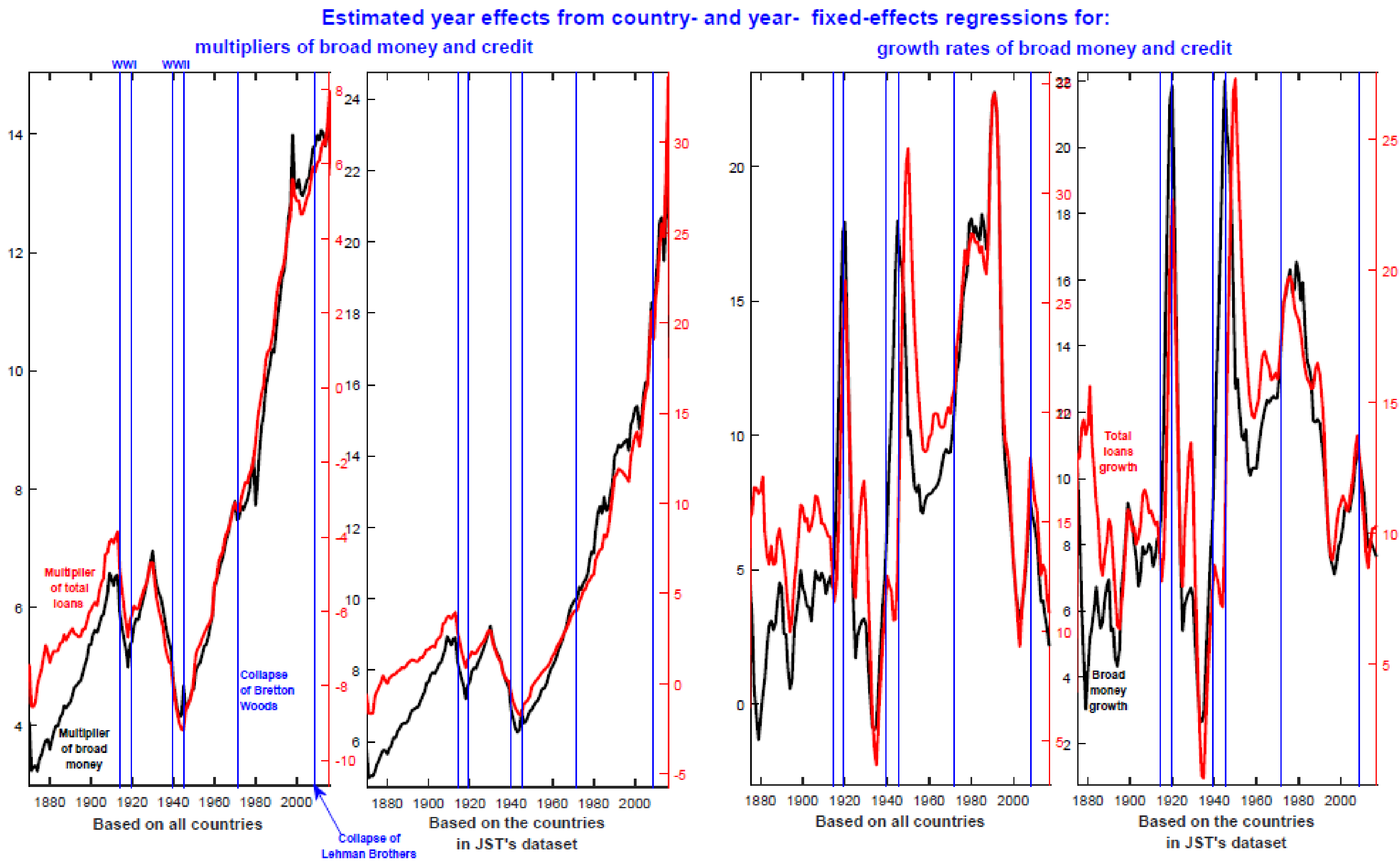

Figure 2 Estimated year effects from country- and year- fixed-effects regressions for either the multipliers, or the growth rates, of total loans and broad money (1870-2017)

Consistent with the evidence in Figures 1a-1c, based on either dataset the estimated year effects for the two multipliers have consistently exhibited a remarkably strong correlation over the entire sample, to the point that, since WWI, they have been nearly indistinguishable.[23] In particular, the *only* apparent change in the relationship between the year effects pertaining to the two multipliers has been an increase in the slope around, or in the aftermath of WWI, with a given increase in μ_M having been associated, until WWII, with comparatively smaller increases in μ_C than thereafter. With this single exception, the relationship between the two multipliers appears to have remained unchanged since 1870.

Although the present work is narrowly focused on the multipliers of broad money and credit, the last two panels of Figures 2 and 3 report the same evidence as in the first two panels of Figure 2, but this time for the growth rates of the two aggregates[24] and, respectively, for the ratios between either aggregate and nominal GDP.

Evidence based on the growth rates points towards a very strong and stable correlation between the year effects for money growth and credit growth since WWI, and a much weaker correlation before that. Evidence based on the ratios with nominal GDP, on the other hand, shows that the correlation between the year effects based on money and credit had been extraordinarily strong up until WWI, and it has also been very strong since the collapse of Bretton Woods. In contrast, the period in between had been characterized by a very long-lasting, but ultimately temporary divergence between the two year effects, which had most likely been associated with the shocks of the two World Wars and the Great Depression. By the early 1970s, however, the divergence had disappeared, and the relationship which had prevailed up until WWI had ultimately reasserted itself.

For the present purposes, the crucial point to stress is that based on either the multipliers, the growth rates, or the ratios with GDP, evidence clearly and consistently suggests that at least since WWII (and likely since WWI), the year effects based on broad money and credit have uniformly exhibited a very strong and stable relationship, often moving essentially in lockstep. This is in contrast with the view, which is dominant within the macroeconomics profession, that since WWII credit has become progressively disconnected from broad money.

The first panel of Figure 3 reports the type of evidence that is typically adduced in support of such a notion. The panel reports, based on either the narrow or the larger dataset, the estimated year effect from country-and-year fixed-effects regressions for the ratio between credit and broad money.[25] Based on either dataset, the estimated

[23] Once again, once appropriately rescaled.

[24] Computed as log-differences.

[25] The evidence based on JST's dataset is conceptually the same as that reported in the bottom panel of Figure 2 of Schularick and Taylor (2012, p. 1035). The only differences are that (i) whereas Schularick and Taylor worked with the logarithms of the ratios, I work with the levels; (ii) Schularick and Taylor considered Denmark, which I instead do not consider for the reason discussed in Section 2; (iii) the latest version of JST's dataset also features Finland and Portugal, which Schularick and Taylor did not analyze.

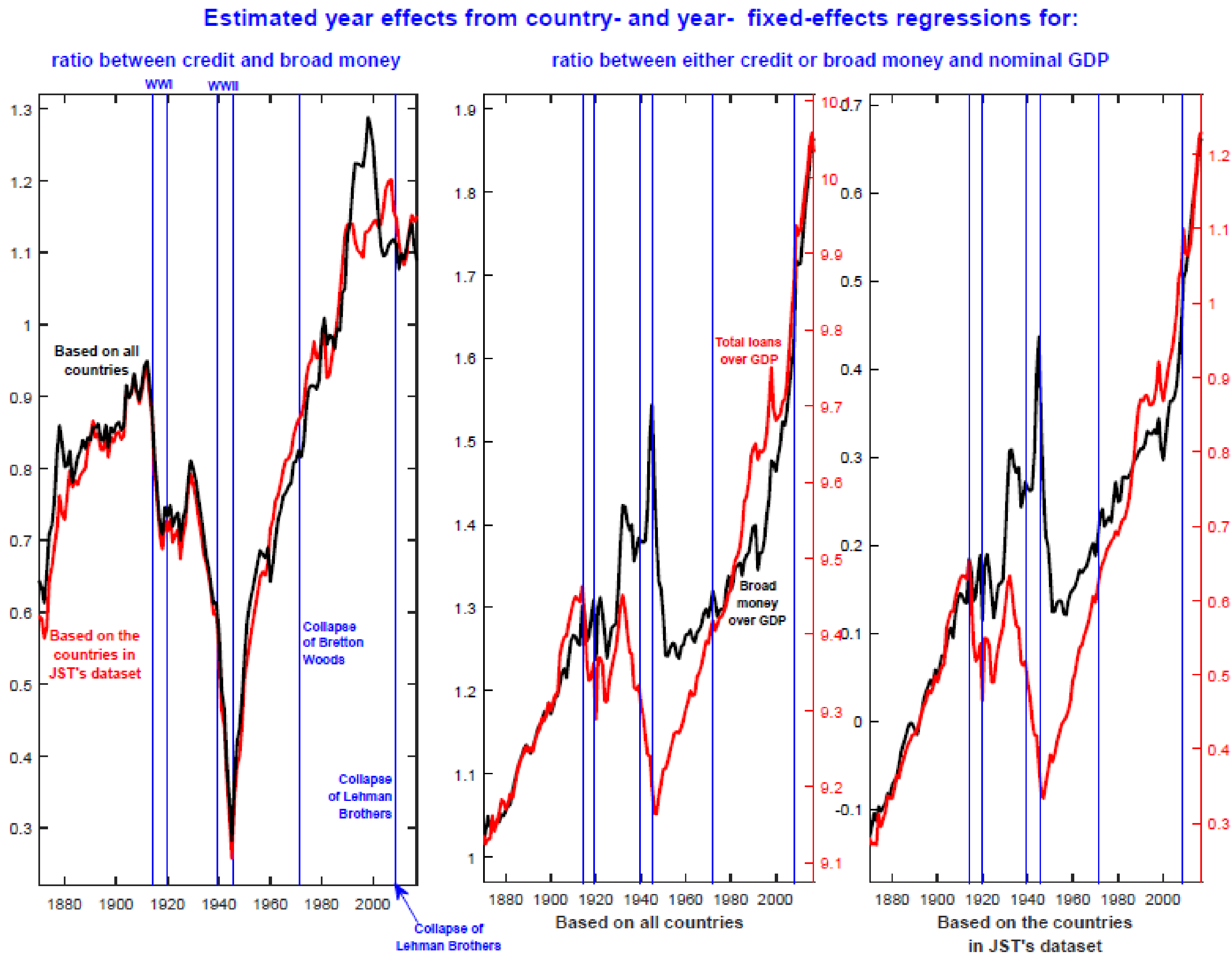

Figure 3 Estimated year effects from country- and year- fixed-effects regressions for the ratio between total loans and broad money, and the ratios between either total loans or broad money and nominal GDP

year effect had exhibited no clear trend until the mid-1930s; it then experienced a dramatic collapse until the end of WWII, and a subsequent rebound which, by the mid-1960s, had returned it at about the levels that had prevailed until WWI; and it has further increased since the collapse of Bretton Woods, although evidence suggests that it may have somewhat stabilized in recent years. For the reasons discussed in the Introduction, however, this evidence is uninformative about the issue of whether since WWII (or over any other period) credit may have become disconnected from broad money.

3.2.1 The scaling variable matters

A comparison between the results for the multipliers in Figure 2, and those for the ratios with GDP in Figure 3, highlights a crucial point: the choice of the scaling variable does matter. The fact that a stable relationship over the entire period since (at least) WWI is identified only when credit and broad money are scaled by the monetary base naturally suggests that the stability uncovered in the present work strictly pertains to the workings of the monetary and financial systems of capitalist economies. On the other hand, the divergence between the two year effects pertaining to the ratios with GDP during the period between the outbreak of WWI and the and of the 1960s suggests that the shocks associated with the two World Wars and the Great Depression had been sufficiently violent to disrupt the relationship between the two aggregates and GDP. At the same time, however, those shocks had *no* discernible impact on the relationship between the year effects extracted from the multipliers, for which the period between 1914 and the end of the 1960s appears in no way different from subsequent years.

3.2.2 Why are multipliers special?

The contrast between the results for the multipliers and those for the ratios between the two aggregates and GDP suggest that, in fact, there is something special about the multipliers. What could this be? As discussed previously, since the monetary base is under the complete control of the central bank, the multipliers of broad money and credit characterize an economy's 'technology' for the 'production' of broad money and credit starting from a given 'input' of base money provided by the central bank. What my evidence shows is that the very nature of such 'technology' is such that, under dramatically different monetary arrangements and macroeconomic conditions, the production of the two aggregates starting from a given input of base money has consistently proceeded in lockstep.

I now turn to evidence from low-frequency regressions, which allow to properly characterize the strength of the relationship between the two multipliers at the low frequencies based on minimal assumptions.

3.3 Evidence from low-frequency regressions

The methodology proposed by Müller and Watson (2018), which is conceptually related to Engle's (1974) band spectrum regression estimator, is based on the notion of extracting low-frequency information from a series of interest by projecting it onto a set of low-frequency periodic functions. In brief,[26] let x_t and y_t, t = 1, 2, 3, ..., T be two series of interest, with $x = [x_1, x_2, ..., x_T]'$ and $y = [y_1, y_2, ..., y_T]'$. Let also $\psi_j(s) = \sqrt{2}\cos(js\pi)$ be a periodic function with period $2/j$; let $\psi(s) = [\psi_1(s), \psi_2(s), ..., \psi_T(s)]'$ be a vector of these functions with periods from 2 to $2/q$;[27] and let Ψ_T be the $T \times q$ matrix whose t-th row is given by $\psi((t\text{-}1/2)/T)$, so that the j-th column of Ψ_T has period $2T/j$. The projections of x_t and y_t onto $\psi((t\text{-}1/2)/T)$ produce the fitted values (i.e., estimated low-frequency components)

$$\hat{x}_t = X_T'\psi((t-1/2)/T) \tag{2}$$

$$\hat{y}_t = Y_T'\psi((t-1/2)/T) \tag{3}$$

for t = 1, 2, 3, ..., T, where $X_T = T^{-1}\Psi_T'x$ and $Y_T = T^{-1}\Psi_T'y$ are the 'cosine transforms' (i.e., cosine-weighted averages of the data) of x_t and y_t. The coefficient in the low-frequency regression of y_t on x_t is

$$\hat{\beta}_{LF} = \arg\min_b E\left[T^{-1}\sum_{t=1}^{T}(\hat{y}_t - b\hat{x}_t)^2\right] = \arg\min_b E\left[\sum_{j=1}^{q}(Y_{j,T} - bX_{j,T})^2\right] \tag{4}$$

The intuition behind Müller and Watson's approach is straightforward. First, the low-frequency components of x_t and y_t are computed by regressing the two series onto a set of low-frequency periodic functions. Then, in a second stage, $\hat{y}_t$ is regressed on $\hat{x}_t$, thus obtaining the low-frequency regression coefficient $\hat{\beta}_{LF}$.

A crucial feature of Müller and Watson's approach is that it has been specifically designed to work well with series characterized by a wide array of low-frequency behaviour, from I(0) processes to (near) unit roots, to cointegrated processes. In what follows I work with the (A, B, c, d) model discussed in Section 3.2.1 of Müller and Watson (2018). The parameterization characterizing this model produces a local-to-zero spectrum of the form

$$S(\omega) \propto A\begin{bmatrix} (\omega^2 + c_1^2)^{-d_1} & 0 \\ 0 & (\omega^2 + c_2^2)^{-d_2} \end{bmatrix} A' + BB' \tag{5}$$

with A and B being (2×2) matrices, with A unrestricted and B lower triangular. As discussed by Müller and Watson (2018, pp. 785-786), the primary motivation behind

[26] For details see Section 2 of Müller and Watson (2018).

[27] For example, for the U.S. the full sample period based on quarterly data, 1875Q2-2019Q4, features T=579 observations, so that by setting (e.g.) q=12 we capture cycles slower than $(2T/q)/4$ = 24.125 years.

Table 1 Evidence from Müller and Watson's low-frequency regressions of μ_C on μ_M

Country	*Period*	Median estimate of ρ and 5-95 per cent credible set	Pseudo p-value[a]
Argentina	1863-1891	0.939 [-0.073 0.989]	0.9405
	1905-2019	0.887 [0.653 0.963]	0.9861
Australia	1976-2017	0.963 [0.750 0.989]	0.4872
Barbados	1990-2018	0.940 [-0.073 to 0.989]	0.8988
Canada	1874-2017	0.918 [0.789 0.963]	0.5415
Chile	1986-2019	0.940 [0.273 0.987]	0.3056
China	1990Q1-2019Q4	0.939 [0.282 0.986]	0.0035
Ecuador	1990-2019	0.937 [0.273 0.986]	0.0524
Finland	1870-1985	0.973 [0.917 0.990]	0.4010
France	1946-1994	0.813 [0.050 0.973]	0.1005
Germany	1883-1913	0.930 [0.246 0.986]	0.9995
	1960-1993	0.911 [0.150 0.984]	0.2030
India	1951-2019	0.925 [0.637 0.977]	0.2776
Indonesia	1985-2019	0.559 [-0.319 0.946]	0.1097
Italy	1870-1997	0.959 [0.889 0.981]	0.7180
Japan	1874-1938	0.635 [-0.001 0.925]	0.9603
	1946-2017	0.943 [0.714 0.981]	0.9991
Malaysia	1975-2019	0.938 [0.497 0.985]	0.1841
Netherlands	1946-1992	0.709 [-0.026 0.960]	0.2577
New Zealand	1960-2016	0.963 [0.841 0.988]	0.4161
Norway	1870-2017	0.962 [0.896 0.985]	0.7105
Paraguay	1990-2019	0.848 [0.001 0.974]	0.2877
Peru	1960-2019	0.704 [0.102 0.925]	0.6926
Portugal	1870-1903	0.199 [-0.570 0.822]	0.4494
	1920-1998	0.829 [0.321 0.963]	0.0095
Saudi Arabia	1980-2019	0.956 [0.639 0.988]	0.2758
Singapore	1991Q1-2019Q4	0.813 [-0.304 0.981]	0.9385
South Africa	1965-2019	0.896 [0.443 0.974]	0.2430
South Korea	1971-2019	0.903 [0.362 0.977]	0.5379
Spain	1900-1935	0.899 [0.124 0.984]	0.7720
	1946-1997	0.971 [0.876 0.990]	0.2188
Sweden	1871-2016	0.952 [0.853 0.975]	0.1410
Switzerland	1907-1950	0.947 [0.571 0.987]	0.2420
	1950-2006	0.963 [0.841 0.989]	0.2365
Taiwan	1962-2017	0.932 [0.604 0.981]	0.3870
Thailand	1976-2019	0.911 [0.371 0.978]	0.6150
United Kingdom	1880-2007	0.981 [0.950 0.990]	0.5529
United States	1880-2007	0.918 [0.841 0.971]	0.8547

For details, see text. [a] Fraction of Monte Carlo replications for which the median estimate of ρ is smaller than that computed in the data, under the null hypothesis that μ_C and μ_M are I(1) and cointegrated.

the (A, B, c, d) model is that it offers a parsimonious, but flexible way of modelling the local-to-zero spectrum, and it comprises, as special cases, several possibilities of interest. For example, $A = 0$ is associated with the I(0) local-to-zero spectrum, whereas $B = 0$, $c = 0$, and $d_1 = d_2 = 1$ yield the I(1) spectrum.

I perform the Bayesian estimation exactly as in Müller and Watson (2018).[28] In what follows I focus on cycles slower than 20 years,[29] and I exclusively consider samples longer than 25 years.

The third column of Table 1 reports the posterior median estimate of the long-horizon correlation coefficient (ρ) in the low-frequency regression of μ_C on μ_M, together with the Bayes equal-tail credible set with 90% coverage.[30] For the sake of simplicity, in what follows I will use 'estimate' as a shorthand for 'posterior median estimate of the long-horizon correlation coefficient'. In order to facilitate the interpretation of the estimates, the fourth column of Table 1 reports a 'pseudo p-value' computed *via* Monte Carlo under the null hypothesis that both μ_C and μ_M are I(1) and they are cointegrated.[31] (The corresponding p-values computed under the alternative assumption considered by Wright (2000), i.e. that the two series are local-to-unity, are qualitatively the same and numerically very close, and they are available upon request.)[32] The pseudo p-value is equal to the fraction of Monte Carlo simulations for which the median estimate of ρ computed based on the simulated data is smaller than the value computed based on the actual data. The intuition, and the reason for considering this metric, is straightforward: if this fraction is very small, this suggests that, at long horizons, the correlation between μ_C and μ_M is significantly smaller than one (i.e., the value corresponding to the case of cointegration). If, on the other hand, this fraction were equal to (e.g.) 50 per cent, this would suggest that the value of ρ computed based on the actual data is compatible with the notion that the long-horizon correlation between the two multipliers is equal to one.

Two main findings emerge from Table 1:

first, in only three cases out of 38[33] (i.e., 7.9 per cent) the pseudo p-value is smaller than 0.1, which throughout the entire paper I take as the benchmark level

[28] I use the MATLAB codes found at Mark Watson's web page

[29] Ideally, I would have preferred to focus on even lower frequencies, but this would have compelled me to discard several post-WWII samples, which are comparatively quite short.

[30] See expression (5) in Müller and Watson (2018). As they point out (see p. 780), ρ^2 is the R^2 in the 'population best linear prediction of the long-run projection $\hat{y}_t$ by the long-run projection $\hat{x}_t$'.

[31] Specifically, I estimate a cointegrated VECM for μ_C and μ_M *via* Johansen's estimator as detailed in Hamilton (1994). I then simulate the estimated data-generation process 10,000 times by bootstrapping the VECM as in Cavaliere *et al.* (2012). To each simulated sample I then apply Müller and Watson (2018) low-frequency estimator, thus obtaining a median estimate of ρ and building up its Monte Carlo distribution.

[32] The Monte Carlo procedure is the same as described in the previous footnote, with the only difference that, instead of bootstrapping the estimated cointegrated VECM, I bootstrap the corresponding near-unit-root VAR in levels. For details, see Benati *et al.* (2021, pp. 55-56 and footnote 24.

[33] China, Ecuador, and Portugal (1920-1998).

of statistical significance. It is important to stress that if (*i*) we were dealing with a 'very large' number of countries, and (*ii*) the DGPs for all of them were orthogonal to one another, the pseudo p-value would be smaller than 0.1 exactly 10 per cent of the time. Although we are here dealing with just 42 countries, whose DGPs cannot strictly speaking be regarded as exactly orthogonal to one another, the same basic logic should approximately hold. This implies that the 7.9 per cent of cases we obtain in which the p-value is smaller than 0.1 are in fact fully compatible with the notion that the long-horizon correlation between the two multipliers is equal to one.

Second, the median estimate of ρ is greater than 0.8 89.2 per cent of the time, and it is greater than 0.9 and 0.95 70.3 and, respectively, 27.0 per cent of the time. This shows that even ignoring the previous argument about statistical significance, and simply focusing instead on the median estimates, in the vast majority of cases the long-horizon correlation between the two multipliers is in fact consistently very high, and close to one. The implication is that even if μ_C and μ_M were not strictly speaking cointegrated, in fact their relationship is, most of the time, very close to that between two cointegrated processes, which at long horizons move one-for-one.[34] In ten cases[35] the median estimate of ρ is greater than 0.95. For each of these countries this simply confirms what the visual evidence in Figures 1*a*-1*c* so starkly suggests, i.e. that at very long horizons μ_C and μ_M move in lockstep, to the point that, once appropriately rescaled, they are often nearly indistinguishable.

3.4 Evidence from cointegration tests

I finally turn to the results from cointegration methods, which I purposefully left for last since they require more stringent assumptions.[36] Tables A.1*a*-A.1*e* in the Online Appendix report bootstrapped p-values[37] for Elliot *et al.* (1996) unit root tests for the multipliers of broad money and credit, either with or without a time trend, and for either the longest available sample, the pre-WWI, post-WWI, and post-WWII sub-samples. The main finding is that only for Argentina the null hypothesis of a unit root is rejected (for both multipliers, and for both the post-WWI and post-WWII periods), whereas for Brazil the evidence is mixed and inconclusive. For all other countries evidence of a rejection of the null of a unit root is either weak or, in

[34] Or, to be precise, one-for-k.

[35] Australia, Finland, Italy, New Zealand, Norway, Saudi Arabia, Spain (1946-1997), Sweden, Switzerland (1950-2006), and the U.K.

[36] For example, Wright's test, that I use in this section, is predicated on the assumption that the series under investigation are either exact or near unit roots. On the other hand (e.g.) country-and-year fixed-effects regressions require no such assumption.

[37] p-values have been computed by bootstrapping 10,000 times estimated ARIMA(p,1,0) processes. In all cases, the bootstrapped processes are of length equal to the series under investigation. As for the lag order, p, since, as it is well known, results from unit root tests may be sensitive to the specific lag order which is being used, for reasons of robustness I consider several alternative lag orders, either 1 or 2 with annual data, and either 2, 4, 6, and 8 with quarterly data.

the overwhelming majority of cases, non-existent.

One possible interpretation of these results is that, with the exception of Argentina, the multipliers of broad money and credit consistently feature an *exact* unit root. Under this interpretation, cointegration tests such as Johansen's (1991) and Shin's (1994), which are predicated on the assumption that the series under investigation are I(1), are indeed appropriate. An alternative, and equally plausible interpretation, however, is that the the two multipliers feature *near* unit roots (i.e., they are 'local-to-unity'). In this case, as shown by Elliot (1998), tests that are predicated on the assumption that the data contain exact unit roots can perform very poorly. Because of this, in what follows I implement the test for cointegration proposed by Wright (2000), which has been designed to be equally valid for data-generation processes (DGPs) featuring either exact or near unit roots. All of the technical details (in particular, the procedure I use to bootstrap the test) are the same as in Benati, Lucas, Nicolini, and Weber (2021, Section 4).

By labelling the normalized cointegration vector between the multipliers of credit and broad money as $[1\ -\beta]'$, Table 2 reports the 90 per cent confidence interval for β based on the longest available samples,[38] whereas Table 3 reports the corresponding evidence for the post-WWI and post-WWII sub-samples. Since I run the cointegrating regression with μ_C as the dependent variable, an estimate of β greater than one, which is obtained in the vast majority of cases, implies that, in the long-run, μ_C increases more then one-for-one with μ_M, thus leading to an increase in the (log) ratio between credit and broad money. This is in line with the evidence in the first two panels of Figure 2, showing that based on either the narrow or the broader dataset, the year effect for the μ_C's has consistently increased faster than the corresponding year effect for the μ_M's. Finally, for the cases in which cointegration is not detected, Table 4 reports the fraction of Monte Carlo simulations for which Wright's test detects cointegration at the 10% level,[39] under the null hypothesis that, in each of these samples, μ_C and μ_M are I(1) and cointegrated.[40]

The main results emerging from the three tables are that,

first, at the 10 per cent level cointegration is not detected in thirteen cases, corresponding to 23.2 per cent of the overall number of samples considered in Tables 2 and 3. In the light of the evidence in Figures 1*a*-1*c*, and in Table 1, most of these instances appear however as puzzling. For example, whereas failure to detect cointegration for Canada for the period 1874-2017 could be rationalized in terms of the disconnect between the two multipliers in the first part of the sample, most of the other rejections are quite difficult to rationalize. This is the case (e.g.) for New Zealand, for

[38] For Argentina, Germany, Japan, Portugal, and Spain data for at least one of the multipliers are not available for the full period dating back to the Gold Standard era. In these cases I report results for two sub-samples for which both series are available.

[39] Corresponding to confidence intervals for β with 90 per cent coverage.

[40] The corresponding evidence obtained under the alternative assumption that μ_C and μ_M are local-to-unity is qualitatively the same, and quantitatively very close to that in Table 4, and it is available upon request.

Table 2 Evidence from Wright's cointegration tests: 90 per cent bootstrapped[a] confidence interval for the second element of the normalized cointegration vector

Country	Period	Confidence interval	Country	Period	Confidence interval
Argentina	1863-1891	[2.9351 31.2500]	New Zealand	1960-2016	NCD
	1905-2019	[0.8280 2.1820]	Norway	1870-2017	[2.2594 3.1017]
Australia	1976-2017	[1.8501 1.9984]	Paraguay	1990-2019	NCD
Barbados	1990-2018	[0.4264 1.0520]	Peru	1960-2019	[5.7904 17.7305]
Canada	1874-2017	NCD	Portugal	1870-1903	[-2.5853 0.6199]
Chile	1986-2019	[0.9578 1.9264]		1920-1998	[0.7913 0.9264]
China	1990Q1-2019Q4	NCD	Saudi Arabia	1980-2019	[1.0102 1.1291]
Ecuador	1990-2019	NCD	Singapore	1991Q1-2019Q4	NCD
Finland	1870-1985	[1.0448 1.1614]	South Africa	1965-2019	[1.2778 1.5569]
France	1946-1994	[1.1228 1.8454]	South Korea	1971-2019	[0.6958 0.8767]
Germany	1883-1913	NCD	Spain	1900-1935	[0.2390 0.4138]
	1960-1993	[1.6340 2.8785]		1946-1997	NCD
India	1951-2019	NCD	Sweden	1871-2016	[-1.6088 0.7255]
Indonesia	1985-2019	[0.4685 1.1114]	Switzerland	1907-1950	[2.2847 2.6205]
Italy	1870-1997	[0.8998 2.1345]		1950-2006	[1.4455 1.5067]
Japan	1874-1938	[0.8590 1.9885]	Taiwan	1962-2017	[0.5187 0.6056]
	1946-2017	NCD	Thailand	1976-2019	NCD
Malaysia	1975-2019	[0.9732 1.1169]	United Kingdom	1880-2007	[1.2755 1.3589]
Netherlands	1946-1992	[1.2050 2.4938]	United States	1880-2007	[0.9686 1.8536]

[a] Based on 10,000 bootstrap replications. For details, see text. NCD = no cointegration detected.

which the two multipliers shown in Figure 1*b* have consistently co-moved closely in synch over the entire sample period. The same holds for China, Ecuador, Germany (1883-1913), India, Japan (1946-2017), Norway (1920-2017), Portugal (1946-1998), and Spain (1946-1997): in all of these cases the visual impression from Figures 1*a*-1*c* clearly suggests that, at long horizons, the two multipliers have co-moved closely in synch. Only for Singapore, in the light of the evidence in Figure 1*c*, failure to detect cointegration does not appear as puzzling. A simple way to rationalize failure to detect cointegration in cases such as New Zealand is in terms of the 'luck of the draw', i.e. as statistical flukes: even if cointegration were there in all samples, due to the very nature of statistical tests, a certain number of failures to detect it ought to be expected.

Table 3 Evidence from Wright's cointegration tests by sub-samples: 90 per cent bootstrapped[a] confidence interval for the second element of the normalized cointegration vector

Country	Period	Confidence interval	Period	Confidence interval
	Post-WWI		*Post-WWII*	
Argentina	1920-2019	[0.3746 1.2082]	1946-2019	[0.3850 2.0323]
Canada	1920-2017	[0.5433 0.5955]	1946-2019	[0.5885 0.6166]
Finland	1920-1985	[0.8626 0.9307]	1946-1985	[0.8284 0.9326]
Italy	1920-1997	[0.8122 1.1609]	1946-1997	[0.8723 1.1369]
Norway	1920-2017	NCD	1946-2017	[0.4552 0.4713]
Portugal	1920-1998	[1.0834 1.2597]	1946-1998	NCD
Sweden	1920-2016	[-0.6805 1.3195]	1946-2016	[-0.7188 1.2812]
United Kingdom	1920-2007	[0.7144 0.7706]	1946-2007	[0.7092 0.7693]
United States	1920-2007	[0.3874 0.7361]	1946-2007	[0.1722 0.7373]

[a] Based on 10,000 bootstrap replications. For details, see text.
NCD = no cointegration detected.

Table 4 Monte Carlo evidence for Wright's tests: fraction of simulations for which cointegration is detected at the 10% level, under the null hypothesis that μ_C on μ_M are I(1) and cointegrated

Canada	1874-2017	0.1170	Norway	1920-2017	0.9379
China	1990Q1-2019Q4	0.2970	Paraguay	1990-2019	0.1700
Ecuador	1990-2019	0.0790	Portugal	1946-1998	0.7989
Germany	1883-1913	0.2290	Singapore	1991Q1-2019Q4	0.6550
India	1951-2019	0.2750	Spain	1946-1997	0.0940
Japan	1946-2017	0.8730	Thailand	1976-2019	0.0840
New Zealand	1960-2016	0.1430			

Second, the Monte Carlo evidence in Table 4 provides some support to this interpretation. For New Zealand, for example, under the null hypothesis that μ_C on μ_M

are I(1) and cointegrated, Wright's test would detect cointegration only 14.3 per cent of the time. By the same token, for Canada, Ecuador, Paraguay, Spain, and Thailand cointegration would be detected between 7.9 and 17.0 per cent of the time, whereas For China, Germany, and India it would be identified between 22.9 and 29.7 per cent of the time. Only for the remaining four countries (Japan, Norway, Portugal, and Singapore) the probability to detect cointegration is quite high, ranging between 65.5 and 93.8 per cent. The implication of the evidence in Table 4 is that, with the exception of the last four countries, failure to detect cointegration on the part of Wright's test should not be regarded as puzzling: rather, the Monte Carlo evidence shows that in all of these cases this is, by far, the most likely occurrence, as the probability to detect cointegration is low, or even very low.

I now turn to discussing several implications of the previous findings.

4 Implications

4.1 Flawed evidence of disconnect between credit and broad money

Since the outbreak of the financial crisis, the evolution of the structure of the financial system over the last several decades has been one of the most intensely investigated issues in macroeconomics. Schularick and Taylor (2012, Figure 2, p. 1035) first documented a sizeable increase in the average ratio between credit and broad money since the end of WWII compared to the previous period, which they have interpreted in terms of a progressive *disconnect* between the two aggregates over the last seven decades:

> 'The first important fact that emerges from the data is the presence of two distinct "eras of finance capitalism" [...]. [T]he first financial era lasted from 1870 to WW2. In this era, money and credit were volatile but over the long run they maintained a roughly stable relationship to each other [...]. Thus, during the first era of finance capitalism, up to 1939, the era studied by canonical monetarists like Friedman and Schwartz, the "money view" of the world looks entirely reasonable. Banks' liabilities were first and foremost monetary, and exhibited a fairly stable relationship to total credit. [...] The relationships changed dramatically in the post-1945 period. [...] [C]redit not only grew strongly relative to GDP, but also relative to broad money after WW2 [...].' [41]

Very similar evidence has subsequently been produced in a series of joint papers with Oscar Jordà.[42]. Jordà *et al.* (2017), for example, speak of a

[41] See Schularick and Taylor (2012, pp. 1034-1036).

[42] See in particular Jordà, Schularick and Taylor (2015, 2017).

'disconnect between credit and (traditionally measured) monetary aggregates'

over the past four decades. Since JST's data only cover the 'traditional' (i.e., non 'market-based', or 'shadow') banking sector, this evidence should be thought of—under such interpretation—as understating the true extent of disconnect between broad money and credit since the end of WWII.[43]

Although this interpretation of the increase in the ratio between credit and broad money since WWII as reflecting a progressive disconnect between the two aggregates has become dominant in the profession, in the Introduction I provided an intuitive explanation for why this evidence is uninformative about the issue at hand. Let us now consider a more formal argument.

Consider a panel of N countries, and assume that, for each country $i = 1, 2, 3, ..., N$, the logarithm of technology in the commercial banking sector, $\gamma_{i,t}$, follows either a random-walk with drift,

$$\ln \gamma_{i,t} = \ln \gamma_{i,t-1} + \delta_i + \eta_{i,t} \tag{6}$$

or a process with a deterministic time trend,

$$\ln \gamma_{i,t} = \kappa_i + \rho_i \ln \gamma_{i,t-1} + \psi_i t + \eta_{i,t} \tag{7}$$

where δ_i is a country-specific drift; κ_i, ρ_i and ψ_i are a country-specific intercept, AR coefficient, and a time trend, respectively; and $\eta_{i,t}$ is a country-specific shock to $\ln \gamma_{i,t}$, with (e.g.) $\eta_{i,t} \sim N(0, \sigma^2_{i,\eta})$. The multiplier of broad money for country i, for the sake of simplicity, is normalized to be equal to technology, i.e. $\ln \mu^M_{i,t} = \ln \gamma_{i,t}$, so that it follows either a random-walk with drift,

$$\ln \mu^M_{i,t} = \ln \mu^M_{i,t-1} + \delta_i + \eta_{i,t} \tag{8}$$

or a process with a deterministic time trend,

$$\ln \mu^M_{i,t} = \kappa_i + \rho_i \ln \mu^M_{i,t-1} + \gamma_i t + \eta_{i,t} \tag{9}$$

where $\mu^M_{i,t} \equiv M_{i,t}/M_{0,i,t}$, with $M_{i,t}$ and $M_{0,i,t}$ being broad money and the monetary base for country i. As documented in Figures 1a-1c, since WWII most (although not all) of the multipliers of broad money have been increasing. I therefore assume that, for all countries, δ_i>0 and ψ_i>0. Finally, I assume that 0<ρ_i<1, so that in (9) $\ln \mu^M_{i,t}$ is stationary around a determistic time trend, whereas κ_i is unrestricted.

Let us then assume, for the sake of the argument, that for each country i the logarithm of the credit multiplier ($\ln \mu^C_{i,t}$) is, up to a stationary stochastic process, a linear transformation of the logarithm of technology, i.e. $\ln \mu^L_{i,t} = \alpha_i + \beta_i \ln \gamma_{i,t} + \theta_{i,t}$, so that

$$\ln \mu^L_{i,t} = \alpha_i + \beta_i \ln \mu^M_{i,t} + \theta_{i,t} \tag{10}$$

[43] This is, e.g., ST's (2012) interpretation: see their footnote 7, p. 1036.

where $\mu^C_{i,t} \equiv C_{i,t}/M_{0,i,t}$, with $C_{i,t}$ being credit (i.e., total loans), and $\theta_{i,t}$ being an I(0) process, e.g., for the sake of the argument, $\theta_{i,t} \sim N(0, \sigma^2_{i,\theta})$. Since the ratio between loans and broad money has been increasing since WWII (see the first panel of Figure 3), I assume that β_i>1 for all i = 1, 2, 3, ..., N.

The implication of (10) is that, conditional on the monetary base, fluctuations in technology, and therefore broad money are the *only* driver of either permanent or long-horizon fluctuations in bank loans (depending on whether the specification for the money multiplier is either (8) or (9), respectively).

From (10), the logarithm of the ratio between credit and broad money is equal to

$$[\ln \mu^C_{i,t} - \ln \mu^M_{i,t}] = \alpha_i + (\beta_i - 1) \ln \mu^M_{i,t} + \theta_{i,t} \tag{11}$$

For the country-and-year fixed-effects regression $[\ln \mu^C_{i,t} - \ln \mu^M_{i,t}] = a_i + b_t + e_{it}$ based on the panel of N countries, the theoretical value of the estimate of the year effect b_t is equal to

$$\hat{b}_t = \frac{1}{N} \sum_{i=1}^{N} (\beta_i - 1) \ln \mu^M_{i,t} \tag{12}$$

which, since β_i>1 for all i = 1, 2, ..., N, exhibits an upward trend originating from the upward trend (either deterministic, or stochastic) in the broad money multipliers. The implication is that this result is not capturing any disconnect between credit and broad money, and an interpretation of this evidence along these lines is therefore incorrect. More generally, the ratio between credit and broad money is uninformative about the issue of whether the two aggregates are, or are not disconnected.

4.2 The true source of the disconnect between credit and broad money

All of the evidence discussed so far pertains to the 'traditional' banking sector, and it shows that, except for commercial banks' greater and greater productivity, the process that, within this sector, jointly creates broad money and credit starting from a unit of monetary base *has not fundamentally changed since the XIX century*. In each country i an increase in log broad money by α_i units (due to an increase in the monetary base and/or to technological progress) has been accompanied, at long horizons, by a corresponding increase in log credit by β_i units, with α_i and β_i having remained constant over the entire sample period.

Drawing from this evidence the conclusion that *nothing* has changed since the Gold Standard era, and in particular since WWII, would however be incorrect. What has changed is what is outside the traditional banking sector which is exclusively featured in either JST's dataset, or the datasets from the Bank for International Settlements and the World Bank, i.e. the *shadow banking sector*. Over the last four decades, shadow banks have experienced an extraordinary expansion. For the present purposes, these institutions feature a crucial difference with respect to the financial

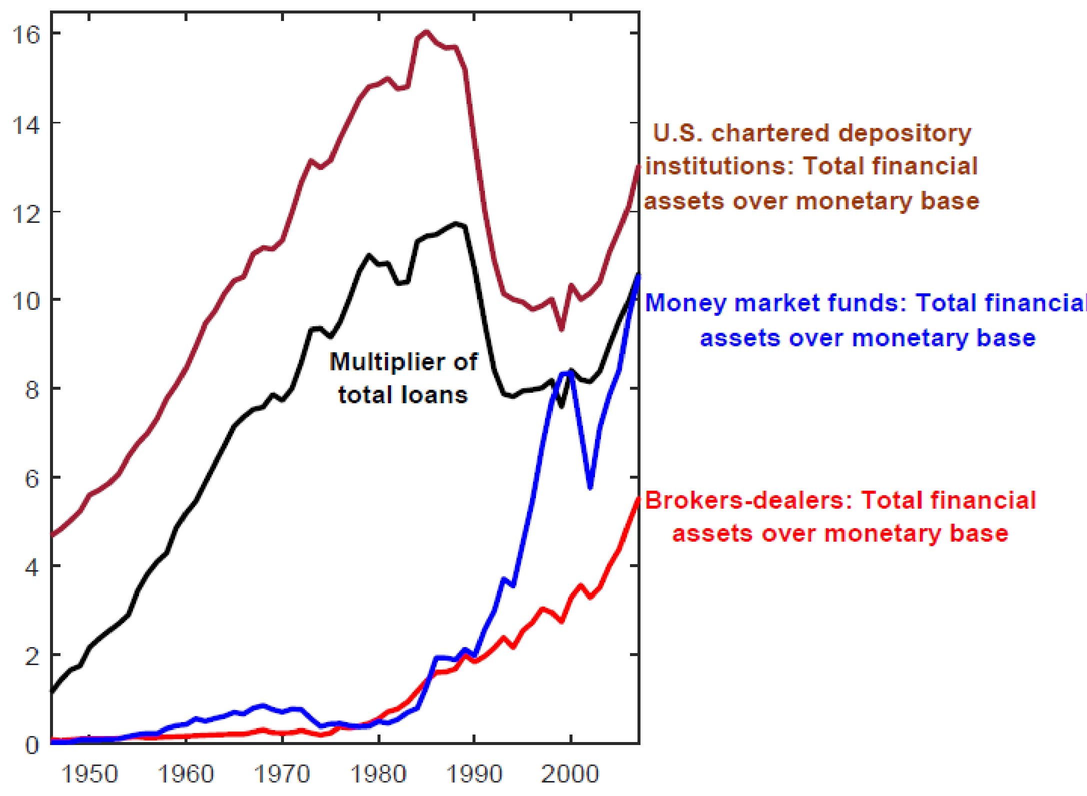

Figure 4 The expansion of the 'market-based' banking system in the United States (1946-2007): ratios between the aggregate assets of financial intermediaries and the monetary base

institutions that belong to the traditional banking sector. Since shadow banks finance the loans they create not by taking deposits from the public, but rather by borrowing on capital markets, such loans feature no corresponding monetary liability that could be counted as part of broad money. As a result, every time shadow banks generate a loan, they automatically introduce a 'wedge' (i.e., a disconnect) between broad money and credit, because they do not create any corresponding monetary liability.

With the notable exception of the U.S., until quite recently data on shadow banks were not being collected systematically, thus making it impossible to perform a comparable analysis for overall credit in the economy (i.e., also including the loans created by shadow banks). In order to gauge an idea about the likely extent of the problem, Figure 4 reports some simple evidence for the U.S. based on annual data from the *Financial Accounts of the United States* (Z.1 release) from the Federal Reserve Board's website. The figure shows, for the period 1946-2007, the ratios between the aggregate assets of several financial intermediaries and the monetary base. I end the sample in 2007 for the reason I discussed in Section 2, pertaining to the dramatic increase in the U.S. monetary base since then. The multiplier of total loans is the same series shown as the red line in the very last panel of Figure 1*c*. Since it pertains to the traditional banking sector, it provides a useful benchmark that allows us to put into proper perspective the *relative* expansion of shadow financial intermediaries compared to traditional banks.

Whereas, as it should be expected based on what we have seen so far, the assets of traditional banks (here, U.S. chartered depository institutions) have closely comoved with the multiplier of total loans, the assets of two market-based intermediaries (brokers-dealers and money-market funds) have skyrocketed since the early 1980s, and they were, as of 2007, of the same order of magnitude and, taken together, even greater than those of traditional banks.

Overall, my evidence suggests that the *only* reason why, today, *overall* credit in the economy (i.e., credit generated by both shadow banks and traditional financial institutions) is partially disconnected from broad money is because of the dramatic expansion of shadow banking over the last four decades. If it were for the traditional banking sector, for which the creation of broad money and credit has proceeded in lockstep since the Gold Standard era, we would still be living in the 'Age of Money', and the 'money view' would still be perfectly relevant. Another way of saying this is that the fundamental distinction is *not* between the two periods before and after WWII, but rather between the traditional and the 'market-based' banking sectors: the former still lives in the 'Age of Money', whereas the ascent of the latter is the only reason why we live in the 'Age of Credit'.

4.3 Implications for macroeconomic analysis

A first obvious implication for macroeconomic analysis is that, when modelling the traditional banking sector, economists should not introduce features creating a 'wedge' between broad money and credit in order to generate a non-existent disconnect between the two aggregates. Rather, they should model a financial sector featuring (at least) (*i*) a traditional banking sector in which the production of broad money and credit starting from a unit of base money proceeds in lockstep at long horizons, and (*ii*) a shadow banking sector that finances the loans it creates by borrowing on capital markets, thus introducing a partial disconnect between *overall* credit in the economy and broad money.

A second implication is the following. Assuming that we have a model featuring (*i*) and (*ii*), the key issue, in my own view, is to come up with a theory for why since WWII, and especially the end of the 1960s, the multipliers of broad money and credit have skyrocketed compared to the previous period (see the first two panels of Figure 2). At first sight, theories of 'monetary over-expansion' under a fiat money regime[44] might be thought to provide an explanation for such pattern. In fact, they do not, as what is needed is not a theory of monetary over-expansion *per se*, but rather a theory of the over-expansion of the money (and credit) multiplier, that is of broad money (and credit) *relative to base money*, rather than in absolute terms. At the same time, it is interesting to notice that the dramatic increases in the multipliers of broad money and credit, compared to their values up until WWII, took place starting from the 1960s, a period characterized, first, by the progressive unravelling of the Bretton Woods system, and then, since August 1971, by the switch towards a pure *fiat* money regime. This timeline could well be purely coincidental, but it could also reflect the fact that, through some mechanism, a *fiat* money regime causes an increase, or even a progressive increase, in the multipliers of broad money and credit.

4.4 Policy implications

The main policy implication is that, by steering broad monetary aggregates, even in the 'Age of Credit' central banks are still perfectly capable of reigning in long-horizon fluctuations in the amount of credit created by the traditional banking sector. On the other hand, since shadow banks feature no monetary liabilities, this strategy has no impact on the amount of loans they create. It is precisely and uniquely because of the loans created by shadow banks that, over the last four decades, overall credit has become partly disconnected from broad money.

5 Conclusions

Money-multiplier analysis played a central role, first and foremost, in Milton Friedman and Anna J. Schwartz's (1963) *Monetary History of the United States*, but over subsequent years and decades it was progressively abandoned, to the point that the money multiplier has almost faded from frontier research, and the very concept has

[44]See Kydland and Prescott (1977) and Barro and Gordon (1983).

largely fallen into oblivion. In this paper I have shown that an analysis of the relationship between the multipliers of broad money and credit sheds a new, and important light on the evolution of the global financial system since the Gold Standard era. My main finding is that, since the XIX century, low-frequency fluctuations in the multipliers of broad money and credit have exhibited an extraordinarily strong correlation in each single one of the 42 countries I analyze. The long-horizon co-movement between the two multipliers has been so strong that, e.g., Wright's (2000) tests consistently detect *cointegration* between the two series, and the long-horizon correlation coefficient in the low-frequency regression of the credit multiplier on the money multiplier produced by Müller and Watson's (2018) methodology is most of the time close to one.

The implication is that, since the XIX century, technological progress has allowed commercial banks to create ever greater amounts of broad money and credit starting from a unit of monetary base. Crucially, however, the *relative* amounts of the two aggregates created out of a unit of base money have remained unchanged over time in all of the countries I analyze.

This finding questions the widespread notion that, since WWII, credit has become disconnected from broad money, and suggests that, except for their greater productivity at creating broad money and credit out of base money, today's commercial banks are not fundamentally different from their XIX century's counterparts. The implication is that *only* the ascent of shadow banks has introduced a disconnect between broad money and credit.

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Online Appendix for: The Joint Dynamics of Money and Credit Multipliers Since the Gold Standard Era

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A The Data

Throughout the paper I report results based on two alternative datasets, a narrow one that comprises 15 countries from Jordà, Schularick, and Taylor's (JST) dataset,[1] and a broader one that also features 26 additional countries.

A.1 The countries in the narrow dataset

For the countries in the narrow dataset, data on *broad money*, total *loans* (i.e., credit), *nominal GDP*, and the *price level* are all from JST's dataset.

As for the *monetary base*, the narrow monetary aggregate featured in JST's dataset (i.e., the series labelled as 'narrowm') is equal to the monetary base only for Norway, Sweden, and the U.S., whereas it is equal to M1 for all other countries.[2] For Norway, Sweden, and the United States I have therefore taken the monetary base from JST's dataset, whereas for the remaining ?? countries I have taken it from national central banks' websites or statistical publications as follows (for Belgium and Denmark, as mentioned in the main text, I was not able to find long-run series for the monetary base). The sources are as follows:

Australia: From the Reserve Bank of Australia's website ('Money base, $ billion, RBA, 42825, DMAMMB').

Canada: From Metcalf, Redish, and Shearer (1998) for the period July 1871-December 1954. After that it is from the *Bank of Canada* ('Monetary base: notes

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[1] The dataset is available from the internet at: http://www.macrohistory.net/data/. The ?? countries are Australia, Canada, Switzerland, Germany, Spain, Finland, France, the United Kingdom, Italy, Japan, the Netherlands, Norway, Portugal, Sweden, and the United States.

[2] See Jordà, Schularick, and Taylor (2021).

and coins in circulation, chartered bank and other Canadian Payments Association members' deposits with the Bank of Canada').

Finland: Table 44.2 ('Monetary Base and Money Stock in Finland. End.of-Morth Data. Millions of Old Marks') of Haavisto (1992).

France: From the Rolnick and Weber (1997) dataset.

Germany: From the Rolnick and Weber (1997) dataset.

Italy: From Fratianni and Spinelli (1997) and, since 1962, from the Table 2 'Componenti della moneta dal 1948 al 1998' of Banca d'Italia, *Tavole Storiche, Indicatori monetari e finanziari*, December 2013.

Japan: From the Rolnick and Weber (1997) dataset until 1970. After that, from the Bank of Japan (''Monetary Base: Average Amounts Outstanding).

Netherlands: From the Rolnick and Weber (1997) dataset.

Portugal: From Table 5 of Mata and Valerio (2011).

Spain: From Barciela-López, Carreras, and Tafunell (2005), Cuadro 9.16, 'Agregados Monetarios', 1865-1998, datos a fin de ano, en millones de pesetas, pp. 697-699.

Switzerland: The two series for the periods 1907-1950, and 1950-2006, are both from the Swiss National Bank's website. Specifically, they both are from the spreadsheet geldmengen.xls. The series for the period 1907-1950 is from the sheet T 1.3, and it is labelled as 'Estimates by Grüebler'. The series for the period 1950-2006 is from the sheet T 1.1 ('Monetary base sources').

United Kingdom: From the Bank of England's spreadsheet of very long-run data millenniumofdata_v3_final.xls, which is available at the Bank of England's website (specifically, the series is in column BL of the sheet A1 ('Headline series').

In those cases in which the original data were available at the monthly frequency, I converted them to the annual frequency by taking annual averages.

A.2 The additional countries in the broader dataset

For the 26 additional countries the sources are as follows.

A.2.1 Loans data

With a few exceptions (discussed below), data on *total credit* provided by *domestic banks* to the *private non-financial sector* are from the *Bank for International Settlements*. The accompanying documentation is available at the *BIS* website.[3] The exceptions are Saudi Arabia before 1992, Peru and Argentina, which are discussed below, and Ecuador, Indonesia, Paraguay, and Taiwan. For these last four countries, I computed nominal total credit to the private non-financial sector based on annual data for nominal GDP (discussed below), and annual series from the World Bank for total credit to the private non-financial sector expressed as a fraction of GDP.

[3] 'Long series on total credit and domestic bank credit to the private non-financial sector: Documentation on data', available at: https://www.bis.org/statistics/totcredit.

As for the other series, the sources for individual countries are the following (when necessary, the series which are originally available at a frequency higher than annual are converted to the annual frequency by taking annual averages).

A.2.2 Other series

Argentina The monetary base ('Base Monetaria, fin de período'), available for the period 1863-2014, is from Table 7.1.2 ('Pasivos Monetarios') from the Banco Central de la República Argentina (the central bank, henceforth, Banco Central). M3 ('M1 + resto de depósitos'), available for the same period, is from Table 7.1.4 ('Agregados Monetarios'). The CPI is from Ferreres (2005) for the period 1900-2004; from INDEC (Instituto Nacional de Estadisticas y Censos) for 2005 and 2007, and from Argentinean Congress for the period 2008 to 2014. Nominal total loans ('Préstamos al Sector Privado') are from Table 38359 ('Préstamos Bancarios') of Ferreres (2005) until 1940, and from the *BIS* after that.[4] Nominal GDP has been reconstructed based on the series for real GDP and the GDP deflator in Table 3.1.1 and 3.3.1 of Ferreres (2005).

Barbados A monthly series for the monetary base, available since January 1990, is from the website of the Central Bank of Barbados, and it has been converted to the annual frequency by taking annual averages. Annual series for Nominal GDP and the CPI, both available since 1975, are from the Barbados Statistical Service. Annual series for the ratio between either broad money, or credit to the private non-financial sector, and GDP are from the World Bank. Corresponding annual series for broad money and credit to the private non-financial sector have been computed as the product of these ratios and nominal GDP.

Brazil Both the monetary base and M2 are from the website of Brazil's central bank. Nominal and real GDP ('valores correntes' and 'valores encadeados a precos de 1995') are from the website of IBGE (the Brazilian Institute of Geography and Statistics). The GDP deflator has been computed as the ratio between nominal and real GDP.

Chile Annual series for nominal GDP and real GDP are from Braun-Llona, Braun-Llona, Briones, Diaz, Luders, and Wagner (1998) for the period 1940-1995. As for the period 1996-2012, they are from the Banco Central de Chile, Chile's central bank (specifically, from the Banco Central's *Anuarios de Cuentas Nacionales*): 'Producto Interno Bruto: Gasto del PIB a precios corrientes, referencia 2013, información histórica (miles de millones de pesos)' and 'Producto Interno Bruto: Gasto del PIB volumen a precios del año anterior encadenado, referencia 2013, información histórica

[4] Over the period of overlapping (1941-2004) the series from the *BIS* dataset is near-identical to that from Ferreres (2005).

(miles de millones de pesos encadenados)'. The GDP deflator has been computed as the ratio between nominal and real GDP. The monetary base and M2 (Monthly averages, billions of pesos) are from Banco Central's *Base Monetaria y Agregados Monetarios Privados*.

China Data on nominal GDP, the CPI, the monetary base, and M2 are from the Federal Reserve Bank of Atlanta's Center for Quantitative Economic Research project on the Chinese economy, at: https://www.frbatlanta.org/cqer/research/china-macroeconomy.aspx.

Colombia The monetary base, M3, nominal GDP and the CPI are from Colombia's central bank's website, at: http://www.banrep.gov.co/en.

Czech Republic The monetary base, M3, nominal GDP and the CPI are from the Czech central bank's website (specifically, the facility called ARAD).

Denmark As I discuss in Section 2 of the main text, for Denmark a series for the monetary base cannot be computed due to the lack of data on commercial banks' *reserves*. Annual series for broad money, total nominal loans, nominal GDP, the CPI, and real GDP *per capita*, all available since 1870, are from JST's dataset.

Ecuador All of the data for Ecuador are from the website of *Banco Central del Ecuador* (henceforth, BCE), Ecuador's central bank. Most of them are from '85 Años, 1927-2012: Series Estadísticas Históricas', a special publication of historical statistics celebrating BCE's 85th anniversary, and they have been updated to 2019 based on data from the BCE. A series for annual CPI inflation is from Chapter 4 of '85 Años'. An annual series for nominal M2 has been constructed by linking the M2 aggregate available for the period 2000-2020 (which is expressed in U.S. dollars, and it has been converted to sucre based on the sucre/dollar nominal exchange rate found in Chapter 2 of '85 Años, and at the website of the BCE), and the M2 aggregate available for the period 1927-1999 (which is expressed in sucre). Both series are from Chapter 1 of '85 Años', and they have been updated based on data from the BCE. An annual series for nominal GDP, available since 1965, is from Chapter 4 of '85 Años'. The series has been updated to 2019 based on data from the CBE. An annual series for the monetary base has been constructed by linking the aggregate available for the period 2000-2020 (which is expressed in dollars and it has been converted into sucre as before), and that from '85 Años', which is expressed in sucre. An annual series for real GDP *per capita*, available since 1947, is from Chapter 4 of '85 Años'. The series has been updated to 2019 based on data from the CBE. Finally, an annual series for credit from domestic banks to the private non-financial sector ('Crédito al sector privado (empresas y hogares)') is from Chapter 1 of '85 Años'. The series is in sucre until 1999, and in dollars since then (converted in sucre as before).

Euro area All Euro area data are from the European Central Bank's website. I wish to thank Alberto Musso for help in tracking down the relevant series' codes.

Greece Base money and M3 are from Boudalis (2016). The CPI is from Greece's central bank ('Table IV: Monthly evolution of the Overall Consumer Price Index during the years 1959-2018'), at Source: http://www.statistics.gr/en/statistics/-/publication/DKT87/.

Hong Kong Quarterly seasonally adjusted series for the monetary base and M3 are from the Hong Kong Monetary Authority (HKMA). The two series are available since 1998Q4 and 1997Q2, respectively. A quarterly seasonally adjusted series for nominal GDP is from Hong Kong's *Census and Statistics Department*. The series has been seasonally adjusted *via* ARIMA X-12.

Hungary Base money, M3 and the CPI are from Hungary's central bank. Nominal GDP ('Gross Domestic Product for Hungary, Millions of National Currency, Quarterly, Seasonally Adjusted') is from FRED II, the St. Louis FED's data portal.

India The monetary base, M3, nominal GDP and the CPI are from India's central bank.

Indonesia Annual series for the monetary base and M2 (available since 1978 and 1985, respectively), and for nominal and real GDP (availble since 1968), are are from Indonesia's central bank. A series for the GDP deflator has been computed as the ratio between nominal and real GDP.

Israel The monetary base and broad money are from Israel's central bank. Nominal and real GDP are from Israel's national statistical agency. The GDP deflator has been computed as the ratio between nominal and real GDP.

Malaysia Base Money ('Jumlah Wang Rizab, Total Reserve Money') is from Table 1.1 from Malaysia's central bank's website. M3 is from Table 1.3.1 from Malaysia's central bank's website. The CPI ('JADUAL 4.1 : INDEKS HARGA PENGGUNA (2000 = 100) MENGIKUT KUMPULAN UTAMA (CHGS), 1980 – 2005, MALAYSIA; Table 4.1 : Consumer Price Index (2000 = 100) by Main Groups (CHGS), 1980 – 2005, Malaysia') is from the central bank's website. Nominal GDP ('KDNK') is from the central bank's website.

New Zealand Series for the monetary base and M3 are from the Reserve Bank of New Zealand's (RBNZ) Long Term Data Series (LTDS) facility at its website. The CPI and nominal GDP are from Statistics New Zealand (New Zealand's national statistical agency).

Paraguay All of the data are from the website of the *Banco Central del Paraguay*, Paraguay's central bank. The monetary base is available since 1990, whereas all other series start between 1960 and 1962.

Peru All of the data are from the website of the *Banco Central de Reserva del Peru*, Peru's central bank. Annual series for the monetary base and M2 are available since 1959. Annual series for nominal and real GDP are available since 1950. A series for the GDP deflator has been constructed as the ratio between nominal and real GDP. An annual series for loans to the private sector is available since 1960.

Poland The monetary base, M3, nominal GDP and the CPI are from Poland's central bank.

Russia Both the monetary base and a broad money aggregate ('Broad money, seasonally adjusted, billions of rubles') are from Russia's central bank. The CPI ('Consumer Price Index: All Items for Russian Federation, Index 2010=100, Quarterly, Not Seasonally Adjusted') is from the OECD's Main Economic Indicators. For nominal GDP I was not able to find a series.

Saudi Arabia Nearly all of the data are from the Excel spreadsheet Annual_Statistics_2020.xls, which is available at the website of the *Saudi Arabia Monetary Authority* (SAMA). Specifically, M2 is from Table 2 of Section 1 ('Money , Banking Statistics and Insurance'), whereas the monetary base has been computed as the ratio between M2 and the M2 multiplier from Table 4 of Section 4 ('Money multipliers'). Total bank credit to the private non-financial sector is from the BIS since 1993. Before that, it is from Table14(a) ('Bank Credit by Economic Activity') of Section 1 (the series has been computed as total credit minus credit to government and credit to financial sector).

Singapore Base money and M3 are from Singapore's central bank's website. The CPI and nominal GDP are from the Department of Statistics Singapore.

South Africa Data for the monetary base, M3, and nominal GDP ('Gross domestic product at market prices, KBP6006J, R millions ') are from the website of the central bank, the South African Reserve Bank (SARB), at: https://www.resbank.co.za. The CPI is from the website of South Africa's statistical agency, at: http://www.statssa.gov.za. The series has been constructed by linking the series labelled as 'Consumer Price Index, VPI00000, all items (The 'general' index) metropolitan' and the series labelled as 'P0141, Consumer Price Index, CPS00000, CPI Headline All urban areas'.

South Korea An annual series for nominal GDP ("Gross domestic product, current prices, Bil.Won") is available from the website of the Central Bank of Korea (henceforth, BOK). Annual series for the monetary base and M2 are from Table 1.1 ('1.1.Money & Banking (Monetary Aggregates, Deposits, Loans & Discounts etc.)') from the BOK's website. The GDP deflator is again from the BOK's website, specifically from Table 10.1.1 ('Main Annual Indicators (reference year 2010)').

Taiwan Annual series for the monetary base ('Reserve money'), M2, and credit are from Taiwan's central bank.

Thailand Base money and M3 are from Table 4 ('Monetary Base (MB), Millions of Baht') and Table 1 ('Financial Survey (M3), Millions of Baht, Liquid liabilities (M3)'), both from Thailand's central bank. The CPI and nominal GDP are also from Thailand's central bank.

A.3 The sample periods

The sample periods for each individual country are shown in Figures 1*a*-1*c*. For each country I consider the longest available sample, with the single exception that, when working with the multipliers, for the United States, the United Kingdom, and the Euro area—whose monetary policies following the financial crisis have led to dramatic expansions in the monetary base—I end the sample period in 2007. The reason for doing so is that including the subsequent period would distort the inference, since the explosion in the monetary base mechanically caused a simultaneous collapse in the two multipliers, thus artificially 'blowing up' the strength of their correlation. On the other hand, for Japan I do not exclude the period of quantitative easing (QE) which started in early 2001, since the expansion in the monetary base was manifestly much more gradual. Also, since within the European Monetary Union (EMU) the monetary base for individual countries is not defined, when I work with money multipliers I necessarily end the samples for these countries in 1998 (at the latest).

Whereas for the countries in JST's dataset the sample periods typically start in 1870,[5] and for Argentina it starts in 1863, for several countries in the broader dataset the samples are quite short. This is the case, e.g., for Brazil, Colombia, and Russia, for which they start around the mid-1990s, whereas for China the sample starts in 1990. In the paper, *all econometric work* for the broader dataset is based on countries whose samples start at least in 1995, whereas I use the countries with shorter samples *only* for 'plotting' purposes, i.e. to visually illustrate the joint dynamics of the two multipliers over the most recent years. Although the samples starting after 1995 are, in fact, quite short, for this paper's purposes the evidence they provide is invaluable, because they clearly show that in countries such as the Czech Republic, the Euro area,

[5] I say 'typically' because, in a few cases, data for the monetary base start much later (e.g., for Australia in 1976) thus compelling me to use shorter samples when working with the multipliers.

Hungary, and Poland, fluctuations in the two aggregates have proceeded in lockstep even in recent years.

As for West Germany, although data on total loans are available since 1947, and data for both the monetary base and broad money are available since 1948, I have chosen to restrict the sample period to 1960-1989. The reason for doing so is the same as in Benati, Lucas, Nicolini, and Weber (2018, see the online Appendix A.12.2). In brief, I am skeptical about the possibility of meaningfully linking the series for the periods 1947 (or 1948)-1959, 1960-1989, and 1990-1998 in order to create continuous series because (*i*) the data before 1960 did not include West Berlin and the Saarland, which, in 1960, jointly accounted for about 6 per cent of overall GDP; and (*ii*) the reunification of 1990 created discontinuities in both GDP and monetary aggregates.

As for Switzerland, as discussed in the text and in the previous Section A.1, the Swiss National Bank provides *two* series for the monetary base, for the periods 1907-1950 and 1950-2006. Unfortunately, the series cannot be linked because their values in 1950 are different: whereas the former series (the series labelled as 'Estimates by Grüebler') is equal, in 1950, to 6267 million Francs, the latter is equal to 5753 million Francs. Therefore, I have decided to consider the two periods separately.

A.4 The dates of the financial crises

For the countries in JST's dataset, the dates of the financial crises are from the dataset itself, and they are the following:

Australia: 1893, 1989.
Canada: 1907.
Switzerland: 1870, 1910, 1931, 1991, 2008.
Germany: 1873, 1891, 1901, 1907, 1931, 2008.
Spain: 1883, 1890, 1913, 1920, 1924, 1931, 1977, 2008.
Finland: 1877, 1900, 1921, 1931, 1991.
France: 1882, 1889, 1930, 2008.
United Kingdom: 1890, 1974, 1991, 2007.
Italy: 1873, 1887, 1893, 1907, 1921, 1930, 1935, 1990, 2008.
Japan: 1871, 1890, 1907, 1920, 1927, 1997.
Netherlands: 1893, 1907, 1921, 1939, 2008.
Norway: 1899, 1922, 1931, 1988.
Portugal: 1890, 1920, 1923, 1931, 2008.
Sweden: 1878, 1907, 1922, 1931, 1991, 2008.
United States: 1873, 1893, 1907, 1929, 1984, 2007.

For the other countries the dates are from either Bordo *et al.* (2001), Cecchetti *et al.* (2009), or Laeven and Valencia (2013). Although, as mentioned in the main text, all econometric work for the broader dataset is based on countries whose samples start at least in 1995, for the sake of completeness in what follows I report the dates of the financial crises for all countries. The dates are the following:

Argentina: 1890, 1891, 1931, 1934, 1980, 1985, 1989, 1995, 2001.
Brazil: No crises in this sample.
Chile: 1976, 1981.
China: No crises in this sample.
Czech Republic: No crises in this sample.
Colombia: 1998.
Ecuador: 1998.
Euro area: 2007.
Greece: No crises in this sample.
Hong Kong: No crises in this sample.
Hungary: No crises in this sample.
India: 1994.
Indonesia: 1997.
Israel: No crises in this sample.
Malaysia: 1985, 1998.
New Zealand: 1987.
Paraguay: No crises in this sample.
Peru: 1983.
Poland: No crises in this sample.
Russia: 1998.
Saudi Arabia: No crises in this sample.
Singapore: No crises in this sample.
South Korea: 1997.
South Africa: 1977, 1985.
Taiwan: 1983.
Thailand: 1983, 1997, 1998.

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Tables for Online Appendix

Table A.1*a* Bootstrapped *p*-values for Elliot, Rothenberg, and Stock unit root tests[a] (without a time trend)

		Multiplier of:			
		broad money		total loans	
Country	*Period*	p=1	p=2	p=1	p=2
Argentina	1863-1891	0.1760	0.4490	0.7155	0.7123
	1920-2019	0.0485	0.0121	0.0218	0.0084
Australia	1976-2017	0.5173	0.4546	0.5799	0.5668
Canada	1874-2017	0.9317	0.8915	0.9923	0.9897
Chile	1986-2019	0.0475	0.1503	0.5753	0.5614
Ecuador	1990-2019	0.3419	0.1879	0.4833	0.1984
Finland	1870-1985	0.9324	0.9426	0.9782	0.9870
France	1946-1994	0.9012	0.9048	0.9649	0.9243
Germany	1883-1913	0.6399	0.6186	0.3058	0.1972
	1960-1993	0.3933	0.5249	0.2254	0.3577
India	1951-2019	0.7887	0.8191	0.8698	0.8995
Indonesia	1985-2019	0.3247	0.2808	0.5592	0.4453
Italy	1870-1997	0.9154	0.7877	0.9891	0.9597
Japan	1874-1938	0.9187	0.9455	0.4219	0.4708
	1946-2017	0.9084	0.6442	0.8094	0.5382
Malaysia	1975-2019	0.4751	0.5118	0.3784	0.3644
Netherlands	1946-1992	0.9769	0.9863	0.9005	0.8739
New Zealand	1960-2016	0.4672	0.4911	0.6135	0.6379
Norway	1870-2017	0.9104	0.9740	0.9949	0.9991
Paraguay	1990-2019	0.7137	0.7804	0.6137	0.6373
Peru	1960-2019	0.8524	0.8539	0.6378	0.4355
Portugal	1870-1903	0.3423	0.4345	0.1011	0.2587
	1920-1998	0.7869	0.7108	0.8746	0.8040
Saudi Arabia	1980-2019	0.6712	0.7463	0.7432	0.6897
South Africa	1965-2019	0.2463	0.2287	0.3691	0.2420
South Korea	1971-2019	0.3915	0.6323	0.3364	0.5080
Spain	1900-1935	0.7966	0.8383	0.3870	0.4618
	1946-1997	0.4746	0.4612	0.4288	0.4282
Sweden	1871-2016	0.9627	0.9332	0.0000	0.5885
Switzerland	1907-1950	0.5436	0.4101	0.5186	0.3460
	1950-2006	0.7188	0.7376	0.6165	0.7761
Taiwan	1962-2017	0.5927	0.6866	0.5588	0.6170
Thailand	1976-2019	0.1453	0.1887	0.2645	0.2800
United Kingdom	1880-2016	0.9488	0.8900	0.9564	0.8857
United States	1880-2017	0.3555	0.3865	0.6226	0.5812

[a] Based on 10,000 bootstrap replications. For details see text.

Table A.1*b* Bootstrapped *p*-values for Elliot, Rothenberg, and Stock unit root tests[a] (with a time trend)

		Multiplier of:			
		broad money		total loans	
Country	*Period*	p=1	p=2	p=1	p=2
Argentina	1863-1891	0.5026	0.8277	0.9703	0.9594
	1920-2019	0.0593	0.0338	0.0602	0.0350
Australia	1976-2017	0.7344	0.6620	0.8692	0.8288
Canada	1874-2017	0.9530	0.9297	0.9447	0.9626
Chile	1986-2019	0.4051	0.6809	0.9019	0.9598
Ecuador	1990-2019	0.2425	0.1251	0.2228	0.0723
Finland	1870-1985	0.9273	0.9551	0.9732	0.9872
France	1946-1994	0.4244	0.5190	0.7595	0.5816
Germany	1883-1913	0.0961	0.1665	0.4265	0.6425
	1960-1993	0.4963	0.3290	0.3654	0.3772
India	1951-2019	0.3740	0.5066	0.8290	0.8760
Indonesia	1985-2019	0.1416	0.1381	0.3035	0.3242
Italy	1870-1997	0.8739	0.5797	0.9951	0.9423
Japan	1874-1938	0.2164	0.3484	0.9657	0.9368
	1946-2017	0.6185	0.5028	0.2378	0.2483
Malaysia	1975-2019	0.4949	0.6663	0.3238	0.4748
Netherlands	1946-1992	0.7591	0.8458	0.3872	0.4562
New Zealand	1960-2016	0.8670	0.8977	0.9463	0.9672
Norway	1870-2017	0.9900	0.9996	0.9989	1.0000
Paraguay	1990-2019	0.6995	0.7415	0.8548	0.8708
Peru	1960-2019	0.5943	0.4519	0.7225	0.3836
Portugal	1870-1903	0.0987	0.3332	0.2902	0.6158
	1920-1998	0.2254	0.0856	0.3482	0.2239
Saudi Arabia	1980-2019	0.3810	0.6145	0.3737	0.6243
South Africa	1965-2019	0.2461	0.2341	0.6082	0.4099
South Korea	1971-2019	0.9240	0.9377	0.7392	0.8627
Spain	1900-1935	0.3297	0.4029	0.5281	0.5613
	1946-1997	0.5339	0.5273	0.5733	0.4845
Sweden	1871-2016	0.9983	1.0000	1.000	1.000
Switzerland	1907-1950	0.4169	0.4305	0.6086	0.4906
	1950-2006	0.6404	0.6896	0.6228	0.7872
Taiwan	1962-2017	0.8224	0.7667	0.9111	0.9323
Thailand	1976-2019	0.6793	0.7028	0.3618	0.6819
United Kingdom	1880-2016	0.8803	0.6605	0.7907	0.6185
United States	1880-2017	0.7345	0.6509	0.7862	0.7378

[a] Based on 10,000 bootstrap replications. For details see text.

Table A.1*c* Bootstrapped *p*-values for Elliot, Rothenberg, and Stock unit root tests[a] for the post-WWI period

		Multiplier of:			
		broad money		total loans	
Country	*Period*	p=1	p=2	p=1	p=2
		Without a time trend:			
Argentina	1920-2019	0.0630	0.0194	0.0371	0.0152
Canada	1920-2017	0.8814	0.8171	0.9539	0.9530
Finland	1920-1985	0.9469	0.9248	0.9804	0.9848
Italy	1920-1997	0.8672	0.6823	0.9785	0.9379
Norway	1920-2017	0.9870	0.9993	0.9980	0.9998
Portugal	1920-1998	0.7805	0.7195	0.8717	0.8042
Sweden	1920-2016	0.0000	0.9558	0.0000	0.5811
United Kingdom	1920-2016	0.8984	0.8319	0.8997	0.8350
United States	1920-2017	0.6226	0.6379	0.7897	0.7715
		With a time trend:			
Argentina	1920-2019	0.0871	0.0702	0.0968	0.0679
Canada	1920-2017	0.6686	0.7545	0.3402	0.6635
Finland	1920-1985	0.9539	0.9655	0.9397	0.9538
Italy	1920-1997	0.7493	0.5613	0.9697	0.8816
Norway	1920-2017	0.6848	0.7778	0.7522	0.8773
Portugal	1920-1998	0.2204	0.0788	0.3500	0.2373
Sweden	1920-2016	0.9852	0.9997	0.9745	0.9884
United Kingdom	1920-2016	0.8311	0.5265	0.6203	0.4672
United States	1920-2017	0.8698	0.7468	0.5272	0.4790

[a] Based on 10,000 bootstrap replications. For details see text.

Table A.1*d* Bootstrapped *p*-values for Elliot, Rothenberg, and Stock unit root tests[a] for the post-WWII period

		Multiplier of:			
		broad money		total loans	
Country	*Period*	p=1	p=2	p=1	p=2
		Without a time trend:			
Argentina	1946-2019	0.1285	0.0577	0.0645	0.0349
Canada	1946-2017	0.6738	0.6108	0.6294	0.6271
Finland	1946-1985	0.7392	0.6966	0.8016	0.8035
Italy	1946-1997	0.6561	0.4969	0.8603	0.8231
Norway	1946-2017	0.9050	0.9718	0.9533	0.9845
Portugal	1946-1998	0.5691	0.5156	0.5997	0.5856
Sweden	1946-2016	0.0000	0.6133	0.0000	0.6004
United Kingdom	1946-2016	0.8156	0.7045	0.7712	0.6930
United States	1946-2017	0.1307	0.2084	0.1598	0.2151
		With a time trend:			
Argentina	1946-2019	0.0787	0.0496	0.1431	0.1078
Canada	1946-2017	0.4724	0.4352	0.7376	0.8238
Finland	1946-1985	0.1103	0.1956	0.1797	0.3377
Italy	1946-1997	0.4292	0.1458	0.9937	0.9428
Norway	1946-2017	0.1973	0.6510	0.6013	0.8470
Portugal	1946-1998	0.2654	0.0879	0.5650	0.3347
Sweden	1946-2016	0.0000	1.0000	0.0000	1.0000
United Kingdom	1946-2016	0.8723	0.5852	0.7544	0.4729
United States	1946-2017	0.3117	0.6233	0.4708	0.7443

[a] Based on 10,000 bootstrap replications. For details see text.

Table A.1*e* Bootstrapped *p*-values for Elliot, Rothenberg, and Stock unit root tests[a]

		Multiplier of:							
		broad money				total loans			
Country	*Period*	p=2	p=4	p=6	p=8	p=2	p=4	p=6	p=8
		Without a time trend:							
Brazil	1995Q1-2019Q4	0.8355	0.7833	0.8124	0.7139	0.9145	0.8224	0.7413	0.6249
China	1990Q1-2019Q4	0.9389	0.7405	0.7410	0.6127	0.9892	0.9605	0.9007	0.8833
Russia	1995Q3-2019Q4	0.7005	0.6528	0.7449	0.7263	0.6314	0.5955	0.6615	0.6750
Singapore	1991Q1-2019Q4	0.3392	0.3047	0.3673	0.2960	0.2005	0.1206	0.0836	0.0387
		With a time trend:							
Brazil	1995Q1-2019Q4	0.4601	0.0804	0.1091	0.0016	0.1570	0.0588	0.0523	0.0050
China	1990Q1-2019Q4	0.3409	0.4913	0.4024	0.6390	0.6688	0.4798	0.7338	0.5807
Russia	1995Q3-2019Q4	0.5794	0.4221	0.4166	0.5842	0.3743	0.3494	0.3731	0.6180
Singapore	1991Q1-2019Q4	0.7527	0.6598	0.7885	0.6329	0.6287	0.3671	0.4511	0.1550

[a] Based on 10,000 bootstrap replications. For details see text.

Figures for online Appendix

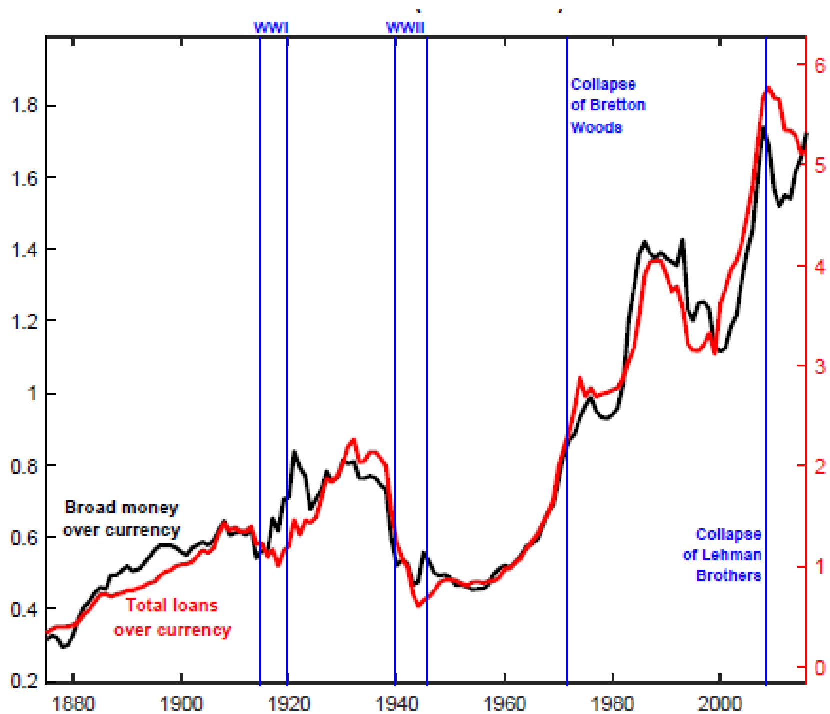

Figure A.1 Denmark (1875-2016): the ratios between either total loans or broad money and currency