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Abstract

This paper reconsiders the long-run demand for M_2 based on a newly constructed dataset featuring 32 countries since the first half of the 19th century. The evidence from cointegration tests suggests that a long-run equilibrium relationship for M_2 demand is hardly present. Specifically, only for five countries (Finland, Korea, Mexico, Paraguay and Taiwan) cointegration tests produce strong evidence in favor of a stable long-run money demand. Evidence for Israel and Lebanon is weaker, but still points towards a stable long-run demand for M_2. For all other countries evidence speaks against a stable money demand or it is mixed across money demand specifications and/or type of cointegration test.

JEL Classification: E4, E41

Keywords: Money Demand, Velocity, Cointegration

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1 Introduction

Economists have been arguing about the usefulness of monetary aggregates for the monetary policy-making process for a very long time. A necessary condition for monetary aggregates to be useful in the conduct of monetary policy is that they are systematically related to real activity, interest rates, and inflation. Over the last three decades, most economists and central bankers have lost confidence in monetary aggregates as information variables for monetary policy. The main reason for the move away from monetary aggregates has been the alleged breakdown of a stable long-run relationship between monetary aggregates, GDP and interest rates. A related reason is the disappearance of the informational content of monetary aggregates for variables such as inflation and nominal GDP within a low-inflation environment (see e.g. Estrella and Mishkin, 1997).

For the United States, for example, there has been a broad consensus in the literature that the stability of the demand for M_1 broke down during the early 1980s,[1] possibly prompted by deregulation and financial innovation. Indeed, recent work by Benati et al. (2018) shows that once they augment the standard M_1 aggregate with either money market deposit accounts (MMDAs) or money market mutual funds (MMMFs) or both,[2] a long-run equilibrium relationship between the velocity of M_1 and the short-term interest rate can be re-established. Benati et al. (2018) further find evidence of a stable long-run demand relationship for M_1 for most of the 31 other countries in their dataset, thus, providing evidence against the widespread view that the long-run money demand is unstable. This new evidence naturally prompts the question: What about broader aggregates?

Broader aggregates better internalize the substitution between different monetary assets that may create instabilities in narrower definitions of money. For example, MMDAs and MMMFs, whose introduction appears to have caused the instabilities in the U.S. M_1 demand, are both included in the standard M_2 aggregate. Until the beginning of the 1990s, U.S. long-run demand for M_2 had indeed been generally regarded as being more stable than that for narrower aggregates (see e.g. Hafer and Jansen, 1991). Since the early 1990s, however, evidence of a stable long-run demand for M_2 has been mixed and inconclusive. One of the most cited studies exploring the stability of U.S. money demand is the work of Friedman and Kuttner (1992). Based on M_2, they documented evidence of a stable money demand relationship for two

[1] see e.g. Friedman and Kuttner (1992).

[2] This adjustment of M_1 was originally proposed by Goldfeld and Sichel (1990).

sub-samples running from 1960 until either 1979 or 1990, but they found no evidence for the period 1970-1990. Following the study of Friedman and Kuttner many authors have extensively analyzed the long-run demand for M_2 for the United States using different samples, money demand specifications and cointegration tests. The resulting evidence is very mixed. Miyao (1996) explores the existence of stable money demand relationships involving M_2 based on several alternative money demand specifications and cointegration tests. In line with Friedman and Kuttner (1992) he finds that evidence of cointegration is stronger for the period 1959-1988, whereas evidence is typically weaker based on sub-samples extending to the early 1990s. For neither period, however, evidence of cointegration is consistent across all of the different tests. Carlson et al. (2000) confirm that the inclusion of data from the 1990s destroys evidence of a stable long-run demand for M_2. In contrast, Lown et al. (2006) find no evidence of a stable M_2 demand for a sub-sample ending in 1988, but they detect a cointegrating relation for two sub-samples including data up to 1994 and 2004, respectively. Haug and Tam (2007) find cointegration in the M_2 demand relationship for a long sample going from 1869 through 1999 and for a post-WWII sample running from 1946 to 1999.

Many studies focus exclusively on the United States[3] and most consider only the post-World War II experience. An exception is the work of Bordo, Jonung (1990, 2009) and Siklos (1993, 1997), who extensively analyze the long-run M_2 demand for the major industrial countries based on a dataset beginning in 1870.

Almost the entire existing literature relies on cointegration tests based on asymptotic critical and p-values. The poor asymptotic properties of cointegration tests are, however, well-known and documented, for example, by Johansen (2002). Further, many studies include ad-hoc variables, such as institutional proxies, break-adjustment dummies or uncertainty measures, in the money demand equation (see, for example, Bordo, Jonung and Siklos), thus making their analyses hard to compare.

In this paper I re-investigate the long-run relationship between M_2, GDP and short-term interest rates based on a dataset featuring 32 countries[4] since the first half of the 19th century, extending the work of Benati et al. (2018) to M_2. For each country I use the longest period for which these three time series are available. To the

[3] Exceptions are documented in Sriram (2000) who provides a survey of empirical money demand studies across a range of countries.

[4] A majority of the countries under investigation are the same as examined by Benati et al. (2018). Since data on M_2 are not available for some countries, I excluded them from the analysis. Instead, a few other countries are added for which data exist on M_2 but not on M_1. Specifically, these countries are Belize, Ecuador, Jordan, Lebanon, Malaysia, Paraguay, Peru and Sweden.

very best of my knowledge, this is the most extensive investigation of the long-run demand for M_2 ever done. Following Benati et al. (2018), this analysis is conducted based on two cointegration tests taking either cointegration or no cointegration as their null hypothesis (Shin's and Johansen's tests, respectively). In order to address the well-known issue of the poor finite-sample performance of cointegration tests, I use bootstrapped p-values, which—as documented *via* Monte Carlo by Benati et al. (2018)—perform better than asymptotic values in small samples. For the countries for which data are available since the 19th century, I perform separate analyses for the Gold Standard[5] and for the subsequent period. The main reason for analyzing the two periods separately is that several key macroeconomic stylized facts have been radically different before and after 1914. In particular, by the very nature of the Gold Standard—which made inflation strongly mean-reverting[6]—based on the Fisher effect nominal interest rates should logically be expected to also be stationary. Since cointegration tests are predicated on the assumption that all of the series under investigation are integrated of order one,[7] both unit root tests and cointegration tests should be performed separately for the two periods.

I consider five alternative money demand specifications. Three of them feature M_2 velocity and a short-term nominal interest rate: they are the traditional semi-log and log-log money demand models, and a specification in the levels of velocity and the short rate along the lines of Selden (1956) and Latané (1960). All these specifications implicitly restrict the income elasticity to one. The other two specifications are unrestricted: one expresses the demand for *nominal* M_2 as a function of nominal GDP and a short-term nominal interest rate, whereas the other models the demand for *real* M_2 as a function of real GDP and an interest rate. Whenever both nominal and real GDP are available, I perform cointegration tests based on all of the five money demand specifications.[8]

The main results I obtain can be summarized as follows. In contrast to Benati et al.'s (2018) results for M_1, for the overwhelming majority of countries cointegration tests do not detect a stable long-run demand for M_2. Specifically, only for five countries

[5] The Gold Standard period is assumed to have ended in 1914, with the outbreak of World War I.

[6] See Barsky (1987) and Benati (2008).

[7] For Johansen's tests see, e.g., Hamilton (1994, p. 571). For Shin's tests, see Shin (1994).

[8] Estimating unrestricted specifications with either nominal or real M_2 and GDP makes sense if and only if the income elasticity is different from one. Otherwise, the two specifications with either nominal or real series are identical. Since income elasticity might differ from one, and since, for many countries, the nominal specification cannot be tested because the series are of a higher order of integration, I consider both real and nominal unrestricted specifications.

(Finland, Korea, Mexico, Paraguay and Taiwan) do tests produce strong evidence of a long-run money demand. Evidence for Israel and Lebanon is weaker, but still points towards a stable long-run demand for M_2. In all other cases evidence is mixed across specifications and/or type of cointegration tests. Further, the log-log specification does not seem to work well for M_2 for any country in the dataset. In fact, it is the only specification under consideration based on which the cointegration tests do not produce consistent evidence in favor of cointegration for any country. This is in contrast to the findings of Benati et al. (2018) for M_1, for which the data seem to prefer the log-log specification over the semi-log and the Selden-Latané specifications for countries in which inflation is very high (e.g. for countries such as Argentina or Israel).

The paper is organized as follows. In the next section, I introduce the newly constructed dataset, and I discuss the visual evidence of the relationship between M_2 velocity and the short rate. In section 3, I discuss the empirical strategy. In sections 4 and 5, I assess the integration properties of the data and the persistence of the cointegration residuals, respectively, whereas in section 6, I present and discuss the results from cointegration tests. Section 7 concludes.

2 A preliminary look at the data

Before introducing the statistical methods, I briefly discuss some noteworthy features of the dataset, and I show visual evidence of the relationship between M_2 velocity and the short rate.

2.1 The data

The dataset features annual data for 32 countries. I select countries with uninterrupted annual time series of nominal M_2, nominal GDP, and a short-term nominal interest rate[9] for at least 30 years. In a few cases (detailed below), data for real GDP and the GDP deflator are also available. The sample length varies across countries: the shortest sample features 30 years (for Germany and Mexico), whereas the longest one (for the United States) features 100 years. A detailed description of the data can be found in appendix D. I use annual data because annual series are typically available for longer time spans than higher-frequency data. Since cointegration is a

[9]For Brazil, Paraguay and Peru no long enough series for the short rate is available. In those cases, I use consumer price inflation or, for Brazil, GDP deflator inflation, as an opportunity cost variable.

long-run concept, the frequency of data is, as a matter of logic, irrelevant for the issue at hand, since increasing the sample size by increasing the sampling frequency (instead of increasing the sample length) does not increase the extent of information for the issue at hand.[10]

For the United States I consider two different series for M_2 for the period 1915-2014. The first series is constructed by linking the M_2 series from Friedman and Schwartz (1970), for the period 1915-1958, to the M2SL series from the St. Louis FED's FRED II for the period 1959-2014. From now on, this series is referred to as the Friedman-Schwartz M_2 aggregate ($FS\ M_2$). The second series is from Lucas and Nicolini (2015), and in what follows is referred to as the Lucas-Nicolini M_2 aggregate ($LN\ M_2$). The main difference between the two series is that for the period 1947-1958 Lucas and Nicolini use data from Robert Rasche (1990), and afterwards they use end-of-period data from FRED. For robustness reasons, in what follows I consider both series. In order to make the overall discussion more concise, however, in the main text I only report and discuss results based on the Friedman-Schwartz aggregate, whereas I mention those based on the Lucas-Nicolini aggregate only if they differ in a material way from those based on the Friedman-Schwartz aggregate.[11]

2.2 Visual evidence on the relationship between M_2 velocity and the short rate

Figures 1-3 show, for all countries in the dataset, M_2 velocity together with a short-term nominal interest rate. In order to better highlight the relationship between the two series, I have subtracted from either of them its sample mean, and I have divided it by its sample standard deviation, thus making them unit-free. For the high-inflation countries (Argentina, Brazil, Chile, Israel, Peru and Venezuela) I show the logarithms of the series, whereas for all other countries I show the levels.

Figures 1 and 2 show the series for the post-WWI period. A strong and positive relationship between M_2 velocity and the short rate is visible for Belize, Finland (1946-1985), Israel, Korea, Lebanon, Mexico and Taiwan. For most of these countries, velocity declines over the entire sample period, thus tracking quite closely the downward trend in the short rate. The pattern differs for Finland and Lebanon, with

[10] See e.g. Siklos (1993); Hendry (1986); Perron (1989).

[11] I focus on the Friedman-Schwartz M_2 aggregate for the following reason. Looking at the log-differences of either M_2 aggregate on the left hand side of figure C.9 one can see that they are essentially identical up to 1947, and reasonably close starting from 1949. In 1948, on the other hand, there is a huge difference between Friedman-Schwartz aggregate and the one from Lucas-Nicolini: in particular, the one from Lucas-Nicolini exhibits a spike, which makes this series less plausible.

the short rate and velocity trending upwards over the entire period in the former case and with both series exhibiting a hump-shaped relationship in the latter case. For Guatemala and Japan a relationship between the short rate and M_2 is visible, but it is less strong. For Argentina, the United States and Turkey a relationship between the two series is only apparent for certain sub-periods. For Argentina a close relationship is apparent until the mid-1950s, whereas since then the two series have behaved differently, with velocity increasing sharply until the beginning of the 1960s, and then declining until the mid-1970s, and the short rate steadily increasing over the entire period until 1984. For the United States the relationship had been strong until the mid-1960s, whereas it has been much less apparent ever since, with the sensitivity of velocity to interest rate fluctuations being smaller than before, and the peaks in the two series being about a decade apart.[12] Turkey's interest rate exhibits a hump-shaped behavior, which is, since the late 1980s, roughly mirrored by M_2 velocity. Strikingly, for a group of countries the relationship between the short rate and M_2 velocity even appears to be negative. This is the case for Canada, Italy, Paraguay, South Africa, Spain and Venezuela. For the remaining countries, the data do not point towards any obvious relationship between M_2 velocity and the short rate.

Figure 3 shows evidence for the Gold Standard period. Interestingly, in almost all cases M_2 velocity follows a downward trend which flattens in the early 1900. The short-term interest rate, however, only follows a similar pattern in the cases of Italy and the United States. In most other cases the short rate exhibits some sort of an inverse hump-shaped pattern.

Overall, visual evidence is very mixed and, in general, it does not point towards a strong and stable long-run relationship between M_2 velocity and the short rate in most countries. The subsequent statistical analysis will confirm that detecting a cointegration relationship between M_2 velocity and the short rate is difficult indeed.

[12] Based on the Lucas-Nicolini M_2 aggregate the relationship is even less apparent, as can be seen in figure C.9 in the appendix.

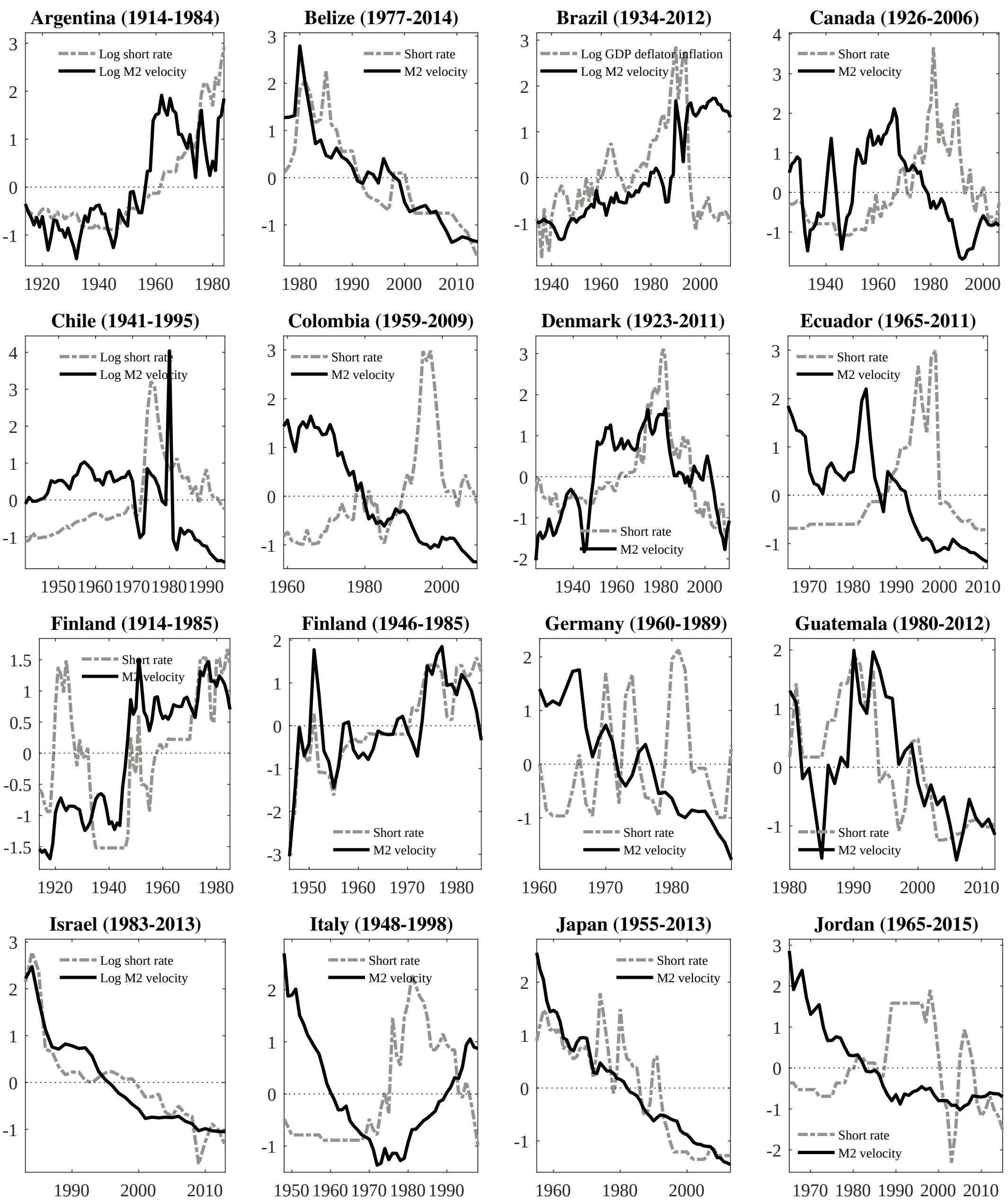

Figure 1: M_2 velocity and the short-term nominal interest rate (all series are demeaned and standardized)

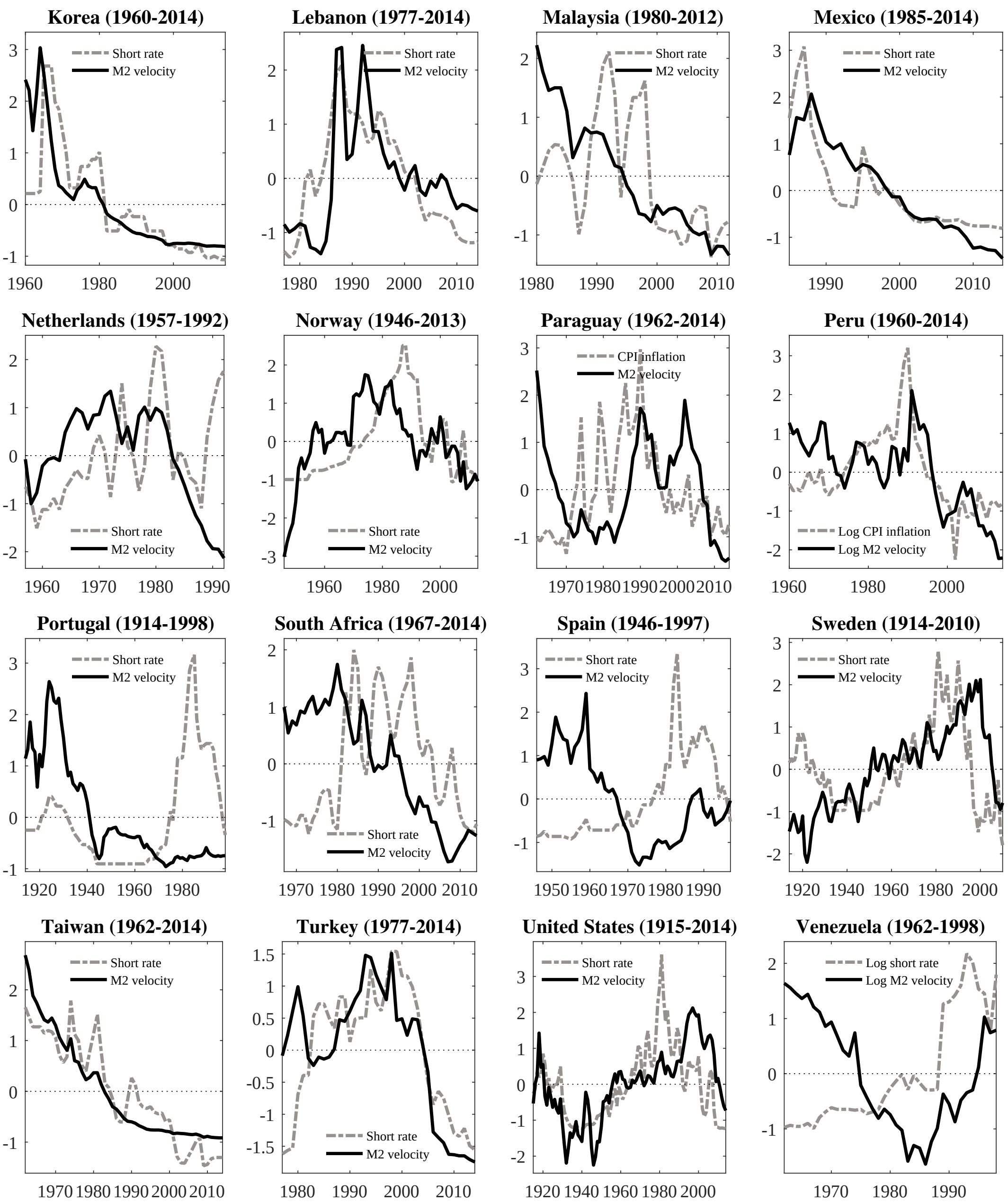

Figure 2: M_2 velocity and the short-term nominal interest rate (all series are demeaned and standardized)

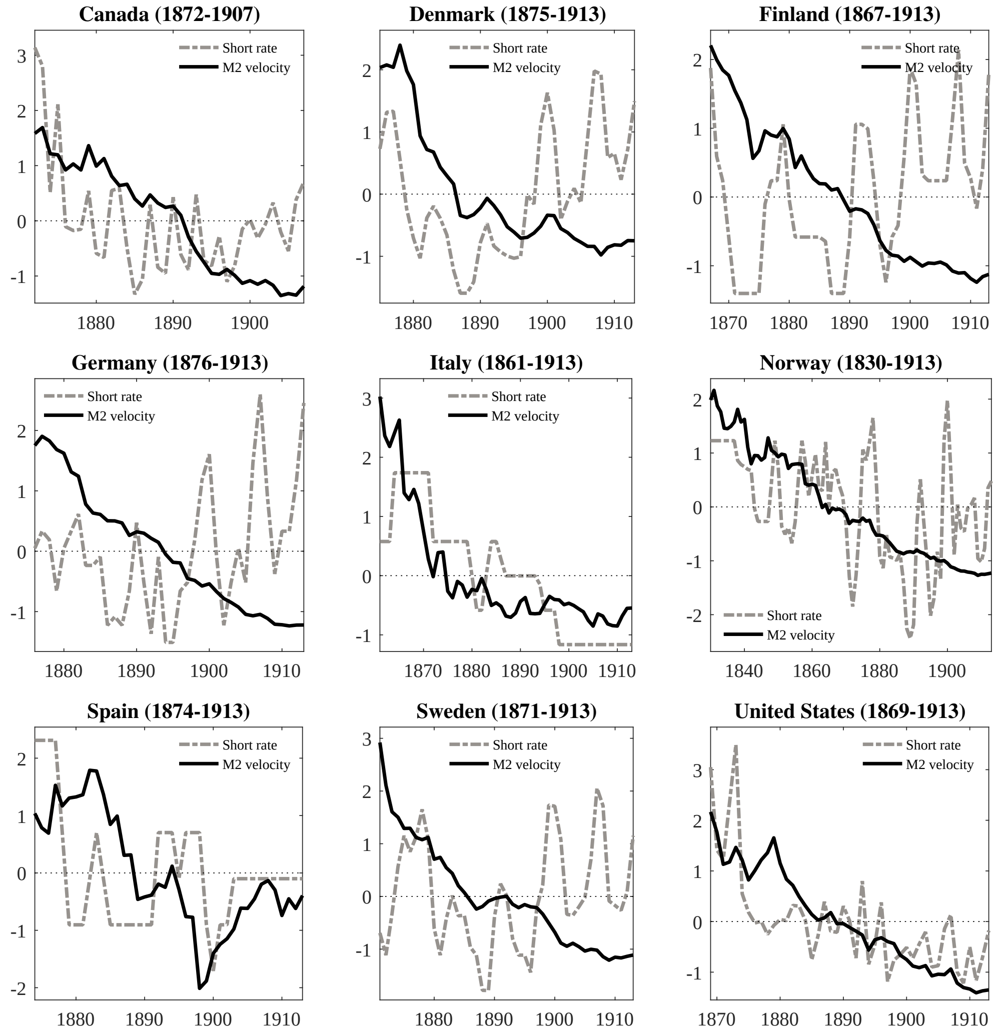

Figure 3: M_2 velocity and the short-term nominal interest rate (all series are demeaned and standardized)

3 Empirical methodology

I explore the long-run demand for M_2 for five different money demand specifications *via* cointegration methods.

3.1 Money demand specifications

In general, the demand for money (M), either nominal or real, is expressed as a function of either nominal or real income (Y) and an opportunity cost variable (R), with money demand being typically assumed to be increasing in income, and decreasing in the opportunity cost:

$$M = \underset{(+),\,(-)}{f(Y, R)}. \tag{1}$$

Another way of expressing money demand is in terms of money velocity (V), defined as the ratio between nominal income and the nominal money stock ($V = Y/M$), with velocity being typically assumed to be an increasing function of the opportunity cost of money:

$$V = f(\underset{(+)}{R}). \tag{2}$$

The analysis in this paper considers five different money demand specifications. I mostly focus on the three bivariate money demand systems, featuring M_2 velocity and a short-term nominal interest rate as opportunity cost variable. Let M, Y, and R denote nominal M_2, nominal output, and a nominal interest rate (with m, y and r denoting the corresponding variables in logarithms). The three bivariate long-run money demand functions can then be written as:

$$\ln\left(\frac{M_t}{Y_t}\right) = m_t - y_t = \mu + \theta_r R_t + \varepsilon_t, \tag{3}$$

$$\ln\left(\frac{M_t}{Y_t}\right) = m_t - y_t = \mu + \theta_r r_t + \varepsilon_t, \tag{4}$$

$$\frac{M_t}{Y_t} = \frac{1}{\mu + \delta R_t} + \varepsilon_t, \tag{5}$$

where θ_r is the interest rate (semi-) elasticity.

Equations (3) and (4) describe the standard 'semi-log' and 'log-log' specifications, which have dominated the literature on long-run money demand since the early 1960s, whereas equation (5) is along the lines of the money demand specifications estimated by Richard Selden (1956) and Henry Allen Latané (1960). This specification postulates a linear relationship between money velocity, defined as the ratio of nominal GDP to nominal M_2, and the short rate: $V_t = \mu + \delta R_t + \varepsilon_t$. In what follows it will be referred to as the Selden-Latané specification. The rationale for considering this specification is that, as shown by Benati et al. (2018), based on M_1 the data seem to prefer it for several low-inflation countries (first and foremost, the United States).[13]

In the bivariate money demand specifications the income elasticity is restricted to one. Since income elasticity might differ from one, I also perform cointegration tests for two unrestricted specifications featuring either nominal or real M_2 and GDP, and the short rate. If, however, the income elasticity actually is equal to one, the two specifications with either nominal or real series are identical and it does not make sense to estimate both specifications. I consider both specifications because, for many countries, the nominal specification cannot be tested because the series are of a higher order of integration. Equations (6) and (7) describe unrestricted money demand specifications featuring the logarithms of either nominal or real M_2 and GDP, and the nominal interest rate in levels.[14]

$$m_t = \mu + \theta_y y_t + \theta_r R_t + \varepsilon_t, \tag{6}$$

$$m_t - p_t = \mu + \theta_y (y_t - p_t) + \theta_r R_t + \varepsilon_t, \tag{7}$$

where p is the price level in logarithms and θ_y is the income elasticity.

3.2 Cointegration tests

I investigate the existence of a long-run demand for M_2 *via* cointegration methods (see Engle and Granger, 1987). Since the variables entering the money demand function are typically found to be non-stationary, cointegration analysis has become the standard method for searching for a stable long-run money demand relationship (see, e.g., Friedman and Kuttner, 1992; Stock and Watson, 1993; Miyao, 1996). In a cointegrated system the long-run relationship between the series is driven by permanent shocks. Permanent changes in one variable map into corresponding permanent changes in the other variables of the system. Further, the finding of a

[13] A theoretical derivation of the Selden-Latané specification can be found in Benati et al. (2018).

[14] For the high-inflation countries – Chile, Israel and Venezuela – the short rate is in logarithms.

cointegration relationship implies that the error term—also known as cointegration residual—of the cointegration equation is stationary. Stationarity of the cointegration residual in turn implies that any deviation from the system's long-run equilibrium is transitory, and thus is bound to disappear in the long-run, so that the equilibrium is restored.

Following Benati et al. (2018), I perform cointegration tests that take either cointegration (Shin's (1994) test) or no cointegration (Johansen's (1988, 1991) tests) as the null hypothesis. Johansen's tests consist of two types, either the trace or the maximum eigenvalue test. The former is a test of the null hypothesis of no cointegration against the alternative of one or more cointegrating vectors, whereas the latter is a test of the null hypothesis of h cointegrating vectors against the alternative of $h+1$ cointegrating vectors, with $h = 0, 1, 2, \ldots$. Shin's test, on the other hand, is a residual-based test of the null hypothesis that the residual of the cointegrating regression has no random walk component. I bootstrap critical and p-values based on the process which is relevant under the null hypothesis, setting the number of bootstrap replications to 10,000.

As for Johansen's tests, I bootstrap trace and the maximum eigenvalue test statistics as in Cavaliere et al. (2012) (henceforth, CRT).[15] Benati et al. show *via* Monte Carlo simulations that if the data-generation process (henceforth, DGP) *does not* feature cointegration, bootstrapping the p-values as in CRT systematically and often significantly improves the performance of the tests in small samples. If, however, a system *does* feature cointegration, bootstrapping Johansen's tests performs well *if and only if* the persistence of the cointegration residual is sufficiently low and/or the sample size is sufficiently large.[16] This is in line with some of the evidence reported in Engle and Granger (1987), who first pointed out how, in small samples, a highly persistent cointegration residual makes it very difficult to detect cointegration based on the Dickey-Fuller test statistic. The intuition is that, in finite samples, a cointegrated process with a highly persistent cointegration residual is difficult to distinguish from a non-cointegrated process. In particular, the more persistent the

[15]For the test of the null hypothesis of no cointegration against the alternative of one or more cointegrating vectors this is a non-cointegrated vector autoregression (VAR) in differences. For the test of the null hypothesis of h versus $h+1$ cointegration vectors this is the VECM estimated under the null of h cointegrating vectors (i.e., for the test of 0 versus 1 cointegration vectors it is a non-cointegrated VAR in differences as for the trace test).

[16]In particular, the Monte Carlo evidence in Benati et al. (2018, Table E.1) shows that for a sample length of $T = 100$ and a persistence of $\rho = 0.75$ ($\rho = 0.9$), the null hypothesis of no cointegration is rejected in 43.3 (16.7) percent of the times. For $T = 50$ and $\rho = 0.75$ ($\rho = 0.9$), the null hypothesis of no cointegration is rejected in 18.4 (11.7) percent of the times.

residual, and the shorter the sample, the more difficult it is to distinguish between the two.

As for Shin's test, Benati et al. (2018) propose a bootstrap procedure which is based on the same general principle underlying CRT's, that is: computing critical and p-values by bootstrapping the process which is relevant under the null hypothesis. Within the present case, the process to be bootstrapped is a vector error-correction model (VECM) estimated under the null hypothesis of one cointegrating vector. Shin's test statistic is computed based on a model with an intercept, but no time trend.[17] Benati et al.'s Monte Carlo simulations show that if the true DGP *does* feature cointegration and the persistence of the cointegration residual is high, the proposed bootstrapping procedure performs better than Shin's asymptotic critical values.[18] However, if the true DGP *does not* feature cointegration, even in large samples cointegration is rejected far too seldom.[19] Lack of rejection of the null of cointegration thus cannot be interpreted as strong evidence that the series are truly cointegrated.

In summary, Johansen's tests detecting cointegration can be taken as strong evidence in favor of cointegration, whereas Shin's tests not rejecting the null of cointegration only represents weak evidence in favor of cointegration. On the other hand, if Johansen's tests do not detect cointegration, this might be due to the high persistence of the cointegration residual and/or a short sample period. Finally, Monte Carlo evidence suggests that, in general, Shin's test is less informative than Johansen's. Therefore, in what follows I will de-emphasize the results produced by Shin's tests, and I will instead mostly focus on those produced by Johansen's methodology.

Benati et al.'s Monte Carlo evidence shows that the performance of either Johansen's or Shin's tests crucially depends on the persistence of the cointegration residual. Before examining the results produced by cointegration tests, I will therefore explore the persistence of the candidate cointegration residuals (defined below) produced by each of the five previously discussed money demand specifications. In the next section I start by discussing the integration properties of the data. For all countries for which

[17] See equation (2) in Shin (1994, p. 93). The test statistic is computed using the Quadratic Spectral kernel. Following Shin (1994) and Benati et al. (2018), the number of leads and lags used in the dynamic OLS estimation of Shin's regression is set to $K = [T^{1/3}]$, where $[x]$ stands for 'the largest integer of x'.

[18] The critical values reported in Shin (1994, p. 100) turn out to be valid only if the cointegration residual is white noise.

[19] In particular, Monte Carlo evidence in Benati et al. (2018, Table E.2) shows that in samples of length T =100, cointegration is rejected at the 10 percent level only in 31 percent of the cases, whereas in samples of length $T = 50$, which is about the length of most of our samples, cointegration is rejected 18 percent of the times.

the I(1) assumption is satisfied for the relevant series, I then explore, in section 5, the persistence of candidate cointegration residuals. In section 6 I then turn to discussing the results produced by cointegration tests, in the light of the respective sample lengths, and of the estimated persistence of the candidate cointegration residuals.

4 Integration properties of the data

A key assumption underlying both Johansen's and Shin's tests is that all of the variables entering either the cointegrated system (in the former case), or the cointegrating regression (in the latter one) are I(1).[20] Before testing for cointegration, the stationarity properties of the series under investigation ought therefore to be ascertained.

Table A.3 in the appendix reports bootstrapped p-values for the Elliott, Rothenberg, and Stock (1996) unit root tests for all series in the dataset. Consistent with the five previously discussed money demand specifications, the tested series are M_2 velocity and the short rate (either in levels or in logarithms), and the logarithms of either nominal or real GDP and M_2. The p-values have been computed based on 10,000 bootstrap replications of estimated ARIMA(p,1,0) processes. For reasons of robustness, for each variable I consider two alternative lag orders, either 1 or 2 years. For the logarithms of either real or nominal GDP and M_2 the estimated models feature an intercept and a time trend. In contrast, for M_2 velocity and the short rate (either in levels or in logarithms), the models include only an intercept, but no time trend.[21] Throughout, I use 10 percent as the benchmark level to test for a unit root. If, for a specific series, the tests based on the two lag order specifications produces contrasting results, I regard the null of a unit root as not having been convincingly rejected, and I therefore assume that the series under consideration has a unit root.[22] In order to ascertain that the order of integration of the series I am

[20] See Hamilton (1994) and Shin (1994), respectively.

[21] As discussed by Hamilton (1994, p. 501) a general principle underlying unit root tests is to choose a specification which *also* represents a plausible description of the data under the alternative hypothesis. The fact that both M_2 and GDP exhibit obvious trends justifies the inclusion of a linear time trend in the estimated model for these series. As for the short rate, on the one hand it is bounded from below by the zero lower bound, whereas on the other hand economic theory suggests that it should not be expected to increase without bounds. As for velocity the exclusion of a time trend is less obvious, since (i) in several countries M_2 velocity exhibits obvious trends, and (ii) Benati et al.'s (2018) conceptual argument for not including a time trend only holds for a demand for money for transaction purposes (that is, for M_1). For these reasons I also test for a unit root in M_2 velocity including a linear time rend, and I report the results in case they differ from those produced by the setup without a time trend.

[22] This is the case for the short rate in levels and in logarithms for Chile (1941-1995), Germany

working with is at most one, unit root tests are also conducted on either their first- or their log-differences.[23] Table A.4 reports the bootstrapped p-values for either the first- or the log-difference of either series, and based on both lag orders. If a unit root can be rejected based on both lag orders, in what follows I assume that the series is I(1). If a unit root is not rejected based on either one or both lag orders, I assume instead that the series' order of integration is greater than one, so that it cannot be included in either a cointegrating system, or a cointegrating regression.

4.1 The short rate

Evidence of a unit root in the short-term nominal interest rate (either in levels or in logarithms) is strong in almost all countries for the period since the beginning of World War I.[24] For the Gold Standard period the tests clearly reject a unit root in the short rate (both in levels and in logarithms) for Norway, Finland and Spain. Unit root tests for the first- or log-differences of the short rate cannot reject a unit root for Morocco and Turkey, thus indicating that either the short rate, or its logarithm, is integrated of an order higher than one. For all of these cases cointegration cannot be tested based on any specification and they are, therefore, excluded from the analysis. For Belize and Italy (1948-1998), the logarithm of the short rate has a order of integration higher than one. As a result, cointegration cannot be tested based on the log-log specification. For Argentina, on the other hand, the short rate in levels has a higher integration order. In this case cointegration cannot be tested based on either the semi-log or the Selden-Latané specification.

4.2 M_2 velocity

Evidence of a unit root in the logarithm of M_2 velocity is strong for all cases except Norway (1946-2013), Israel, and Italy under the Gold Standard: in either of these cases, a unit root can be rejected based on both lag orders. Including a time trend

(1960-1989), Israel, and Canada and the United States under the gold standard, and for the logarithm of the short rate for Mexico, and for the logarithms of velocity for Chile (1941-1995), Finland (1946-1985), Jordan and Taiwan.

[23]For these tests, the estimated models include an intercept, but no time trend.

[24]The only exception is Finland (1946-1985) for which the short rate in logarithms is stationary based on both lag order specifications. Thus, for Finland (1946-1985) the cointegration cannot be tested based on the log-log specification. For Brazil, Paraguay and Peru, for which I either use GDP deflator inflation or CPI inflation as an opportunity cost variable, the series is stationary either in levels (Brazil, Peru) or in logarithms (Paraguay). Thus, for Brazil and Peru cointegration cannot be tested based on the semi-log or the Selden-Latané specification and for Paraguay cointegration cannot be tested based on the log-log specification.

changes the results for Israel, and for Italy under the Gold Standard.[25] In what follows I will therefore assume that either series is non-stationary. For Norway, including a time trend does not produce a different result. With the logarithm of velocity being stationary, it is not possible to perform cointegration tests based on either the semi-log or the log-log specification. In what follows I therefore ignore, for Norway, either of the two specifications. For the level of M_2 a unit root gets rejected more often: either Finland, Italy and Sweden under the Gold Standard; or Chile, Israel, Japan, Jordan,[26] Norway and Taiwan for the period since 1914 exhibit stationary M_2 velocity series. For Finland under the Gold Standard period, for which also the short rate is stationary, the Selden-Latané specification can be estimated using a bivariate VAR in levels.[27] Further, the fact that both series are I(0) is compatible with—although it does represent a proof of—the existence of a long-run demand for M_2. In all of the other cases, on the other hand, it is not possible to perform cointegration tests based on the Selden-Latané specification. Finally, for Italy (1948-1998), Morocco, and Venezuela, and for Canada under the Gold Standard, either the level or the logarithm of velocity is integrated of a higher order than one. In those cases, cointegration cannot be tested based on the relevant specification.

4.3 Nominal GDP and nominal M_2

In most cases, a unit root in nominal GDP and nominal M_2 cannot be rejected at the 10 percent level for either of the two lag orders considered.[28] However, for Canada under the Gold Standard, and for 22 countries[29] for the period since 1914, unit root tests for the variables in log-differences suggest that for nominal GDP and/or nominal M_2 the order of integration is greater than one. These countries will therefore be excluded from the analysis based on unrestricted specifications featuring the logarithms of nominal GDP and M_2.

[25] For Israel the p-values change to 0.185 and 0.001 based on the two lag orders, whereas for Italy (1861-1913) they change to 0.153 and 0.830, respectively.

[26] For Jordan, including a time trend leads to p-values equal to 0.049 and 0.899, respectively. In what follows I will therefore assume that this velocity series is non-stationary.

[27] Estimating a bivariate VAR featuring M_2 velocity and the short rate with 10,000 bootstrap replications yields a bias-corrected median estimate of the coefficient on the short rate of 0.00. The coefficient is not significantly different from zero.

[28] The exceptions are Israel, Mexico and Denmark under the Gold Standard, for all of which nominal M_2 is stationary.

[29] Argentina, Brazil, Canada, Chile (1941-1995), Colombia, Germany, Guatemala, Italy, Japan, Jordan, Korea, Lebanon, Malaysia, Mexico, Morocco, Netherlands, Paraguay, Peru, Spain, Taiwan, Turkey and Venezuela.

4.4 Real GDP and real M_2

The logarithm of real M_2 is uniformly I(1), with the exceptions of Norway during either the Gold Standard or the period since 1914 (in both cases it is stationary); and of Italy (1948-1998), Japan, and Canada during the Gold Standard, for which the order of integration is greater than one.
Real GDP is non-stationary as well, with the exception of Finland (1914-1985), for which unit root tests clearly reject a unit root based on both lag orders. For Canada during the Gold Standard, and for Italy, Japan, the Netherlands, and Taiwan during the period since 1914, unit root tests for the variables in log-differences suggest that the order of integration of log real GDP is greater than one. Therefore, for all of these countries I do not perform cointegration tests based on unrestricted specification including real GDP and real M_2.

4.5 Summing up

Altogether, in 30 cases cointegration can be tested based on the semi-log specification[30] and based on the log-log specification.[31] Cointegration tests based on the Selden-Latané specification can be conducted for 20 countries[32] for the period since 1914, as well as for Denmark, Germany and the United States under the Gold Standard. The unrestricted nominal specification is analyzed for Belize, Denmark, Ecuador, Finland for the period 1946-1985 and for the period 1914-1985, Norway, Portugal, South Africa, Spain, Sweden and the United States since 1914, and for Italy, Sweden and the United States under the Gold Standard. Based on unrestricted specifications featuring real GDP and real M_2, cointegration can be tested for 13 countries[33] for the period since 1914, and for Denmark, Italy, Sweden and the United States under the Gold Standard.

[30] Denmark, Germany, Italy, Sweden and the United States during the Gold Standard; and Belize, Canada, Chile, Colombia, Denmark, Ecuador, Finland for the period 1946-1985 and for the period 1914-1985, Germany, Guatemala, Israel, Japan, Jordan, Korea, Lebanon, Malaysia, Mexico, Netherlands, Paraguay, Portugal, South Africa, Spain, Sweden, Taiwan and the United States over the period since 1914

[31] Argentina, Peru and the countries above, except for Belize and Paraguay

[32] Canada, Colombia, Denmark, Ecuador, Finland for the period 1946-1985 and for the period 1914-1985, Germany, Guatemala, Jordan, Korea, Lebanon, Malaysia, Mexico, Netherlands, Paraguay, Portugal, South Africa, Spain, Sweden and the United States

[33] Canada, Chile, Colombia, Denmark, Finland for the period 1946-1985 and for the period 1914-1985, Israel, Korea, Portugal, South Africa, Spain, Sweden, the United States and Venezuela

5 Assessing the persistence of candidate cointegration residuals

Since, as previously discussed, the performance of cointegration tests crucially hinges on the persistence of the cointegration residual, before turning to the evidence from cointegration analysis I explore the persistence of candidate cointegration residuals. 'Candidate cointegration residual' (henceforth, CCR) is an expression used by Benati et al. (2018) to indicate the linear combination of the I(1) variables in the system which will indeed be regarded as a cointegration residual if cointegration actually is detected.[34]

Tables A.1 and A.2 report Hansen (1999)'s 'grid bootstrap'[35] median-unbiased (henceforth, MUB) estimates of the sum of the AR coefficients in AR(2) representations for the CCRs. From now on, I will refer to this object as $\hat{\rho}_{MUB}$. The results reported in the tables are based on Johansen's estimator of the cointegration vector. For each value of the sum of the AR coefficients in the grid I use 2,000 bootstrap replications.

Table A.1 reports evidence for the bivariate specifications. Based on semi-log specifications, $\hat{\rho}_{MUB}$ ranges from 0.30 for Germany (1960-1989) to 0.98 for Portugal (1914-1998) and Jordan. Based on the log-log specification and the Selden-Latané specification, it lies within a similar range. Across all three specifications, CCRs exhibit a non-negligible, and sometimes high, or very high extent of persistence. Based on all three bivariate specifications, $\hat{\rho}_{MUB}$ is greater than or equal to 0.9 in roughly a third of the cases. Further, in many cases the 90 percent confidence interval includes 1.00. Based on the log-log specification, for eight countries $\hat{\rho}_{MUB}$ is considerably higher than for the semi-log or the Selden-Latané specifications. These countries are Chile, Ecuador, Israel, Japan, Korea, the Netherlands, Spain, and Taiwan.

Table A.2 reports results based on unrestricted trivariate specifications. Out of the 13 possible cases for which cointegration tests can be performed based on specifications featuring nominal GDP and M_2, four have a value of $\hat{\rho}_{MUB}$ exceeding 0.9. For the remaining cases, $\hat{\rho}_{MUB}$ ranges between 0.29 and 0.86. As for specifications featuring real GDP and real M_2, in five cases out of a total of 17 $\hat{\rho}_{MUB}$ is greater than 0.9, in three cases it is between 0.75 and 0.9, and in the remaining nine cases it is below

[34]The reason for labelling it as a 'candidate' is that, as previously discussed, if a cointegration residual is very highly persistent, cointegration will likely *not* be detected.

[35]This 'grid bootstrap' method for the construction of confidence-intervals proposed by Hansen (1999) performs better than conventional bootstrap methods when the sampling distribution depends upon the parameter of interest.

0.75.

Under these circumstances, Johansen's tests will have a hard time detecting cointegration even if it truly is in the data. Based on the Monte Carlo evidence reported in Benati et al. (2018, table C.1), if the true DGP features cointegration, based on a sample period of length $T = 100$ ($T = 50$), and a persistence of the cointegration residual equal to $\rho = 0.9$, Johansen's tests correctly reject the null of no cointegration only 16.7 (11.7) percent of the times. If $\rho = 0.75$, the fraction of rejections is 43.3 (18.4) percent. This suggests that especially in cases characterized by small samples and a highly persistent cointegration residual, Johansen's tests will likely not detect cointegration.

6 Evidence from cointegration tests

This section presents the results from cointegration tests based on any of the five money demand specifications considered herein. I mostly focus on the results based on the bivariate specifications and only briefly comment on the results based on unrestricted specifications.

6.1 Bivariate specifications

Table 1 reports the bootstrapped p-values for Johansen's and Shin's tests based on the semi-log, the log-log, and the Selden-Latané specification. The bootstrapped p-values are based on 10,000 bootstrap replications. Figures B.1-B.6 in the appendix report the estimated CCRs produced by the Johansen procedure for all specifications. More specifically, figures B.1 and B.2 show the estimated CCRs based on the semi-log specification, figures B.3 and B.4, and figures B.5 and B.6 show the estimated CCRs based on the log-log specification and the Selden-Latané specification, respectively. For most countries, either test clearly speaks against cointegration, or they produce contrasting evidence (e.g., with both Johansen's and Shin's tests not rejecting the null). In the light of the visual evidence discussed in section 2.2 this was largely expected. On the other hand, based on the previous discussion of how the persistence of the cointegration residual impacts upon the performance of cointegration tests, we know that even if there is cointegration in the data, given the overall non-negligible persistence in actual CCRs, Johansen's tests will likely have a hard time detecting it. It is therefore essential to carefully interpret the results produced by cointegration tests, in order to determine whether there really is no cointegration relationship in the data, or whether the persistence in the cointegration residual is just too high and/or the sample period too short for the tests to be able to detect cointegration. This requires a careful analysis of the results, taking all necessary features of the data into account.

In general, the semi-log and the Selden-Latané specifications produce very similar results. Based on the log-log specification, on the other hand, Johansen's tests and Shin's test provide contrasting evidence in most cases. In the following, I discuss the individual cases in detail to gain a better understanding of the results.

6.1.1 Cases in which the evidence points towards cointegration

The few cases for which evidence uniformly points towards a cointegration relationship in the data are reported in this sub-section. Based on the semi-log specification, Johansen's and Shin's tests both detect cointegration in the following six cases: Israel, Korea, Mexico, Paraguay, Taiwan, and Italy under the Gold Standard. These results are largely in line with the visual evidence discussed in section 2.2, with the exception of Paraguay for which no apparent relationship between M_2 velocity and CPI inflation is visible. Based on the Selden-Latané specification, the tests confirm the presence of a cointegration relation in the cases of Korea, Mexico and Paraguay. In addition, they detect cointegration for Finland (1946-1985) and Lebanon. For the other three countries the Selden-Latané specification could not be tested due to M_2 velocity being stationary. Finally, based on the log-log specification there is no uniform evidence of cointegration for any country.

6.1.2 Cases in which the tests produce conflicting evidence

Based on the semi-log specification, cointegration tests produce conflicting evidence in 14 cases, with neither Johansen's nor Shin's tests rejecting the null.[36] Specifically, this is the case for Denmark, Germany, Sweden and the United States for the Gold Standard period; and for Belize, Canada, Colombia, Denmark, Ecuador, Finland (1946-1985), Guatemala, Japan, Lebanon, Spain, and the United States based on the Lucas-Nicolini aggregate for the period since 1914. From Benati et al. (2018)'s Monte Carlo evidence on the performance of cointegration tests, which was discussed in section 3, we know that, even in large samples, Shin's test has problems in rejecting the null of cointegration if the true DGP does not feature it. Therefore,

[36] On the other hand, only in a single case do Shin's tests reject cointegration and Johansen's tests reject no cointegration. This is the case for Jordan based on the Selden-Latané specification.

as previously discussed, the fact that Shin's tests do not reject the null does not represent strong evidence that cointegration truly is in the data, and, accordingly, we should not put too much emphasis on these results. Since, on the other hand, Johansen's tests have problems in detecting cointegration if the cointegration residual is highly persistent and/or the sample period is short, I turn to analyzing the persistence and stationarity of CCRs. CCRs appear to be stationary for Denmark, Germany, Sweden and the United States under the Gold Standard; and for Belize, Colombia, Denmark (1923-2011), Finland (1946-1985), Guatemala, Japan, Lebanon, and the United States based on the Lucas-Nicolini aggregate for the period since 1914. For the United States (1915-2014), Denmark (1923-2011), Colombia, Guatemala and Lebanon estimated $\hat{\rho}_{MUB}$'s are quite high (at 0.90, 0.91, 0.88, 0.82 and 0.81, respectively), which could explain why Johansen's tests do not detect cointegration despite the stationary-looking residuals. For Belize, Finland (1946-1985), Japan and the three countries under the Gold Standard, CCRs are low to moderately persistent, but, considering the rather short sample, there is the possibility that Johansen's tests' failure to reject no cointegration is spurious. This is less of a possibility for Canada, Ecuador and Spain, whose CCRs do not appear to be stationary: for these three countries, it can therefore be stated with reasonable confidence that cointegration truly is not in the data.

Based on the Selden-Latané specification results are overall very similar. Evidence is mixed for Belize, Canada, Colombia, Denmark, Ecuador, Guatemala, Spain, the United States based on the Lucas-Nicolini aggregate, for the period since 1914; and for Denmark and the United States during the Gold Standard. The estimated persistence of CCRs, and their stationarity appearance, suggest that cointegration might be possible for Belize, Colombia, Denmark (1923-2011), Guatemala, and the United States under both periods, and Denmark under the Gold Standard. On the other hand, this is less likely to be the case for Canada, Ecuador and Spain.

Finally, based on the log-log specification Johansen's and Shin's tests produce conflicting evidence in 22 cases. In contrast to Benati et al. (2018), who find that, for M_1, the log-log specification works well for several countries—first and foremost, high-inflation ones such as Argentina and Israel—for M_2 this specification does not seem to work for any country. The strong look of non-stationarity of most of the CCRs in figures B.3 and B.4 further adds to this impression: this is especially clear for Argentina, Denmark, Ecuador, Finland, Israel, Japan, Korea, Portugal, South Africa, Spain, Sweden, the United States and Taiwan, and for the United States under the Gold Standard. In all of these cases it is therefore reasonable to argue that

Johansen's tests have correctly captured the lack of cointegration.

6.1.3 Cases in which the evidence is against cointegration

Based on the semi-log specification, for nine countries both cointegration tests point towards no cointegration, with Johansen's tests not rejecting the null of no cointegration, and Shin's tests instead rejecting the null of cointegration. This is the case for the United States (1915-2014),[37] Chile (1941-1995), Finland (1914-1985), Jordan, Malaysia, the Netherlands, Portugal (1914-1998), South Africa, and Sweden (1914-2010). In order to assess whether these results might still be compatible with the presence of cointegration, I next examine the CCRs in figures B.1 and B.2. For Chile, Finland, Jordan, Portugal, and Sweden, they clearly appear non-stationary. For Finland, Jordan, Portugal, and Sweden, $\hat{\rho}_{MUB}$ is indeed very high, ranging between 0.93 and 0.98, and with the 90 percent confidence intervals stretching up to 1.01 in all four cases. For Chile persistence is somewhat lower, with $\hat{\rho}_{MUB}$ being equal to 0.68. Taken together with the visual evidence discussed in section 2.2, absence of cointegration is a reasonable interpretation of the overall evidence. For Malaysia, the Netherlands, South Africa and the United States, on the other hand, the CCRs appear broadly stationary. At the same time, they appear to be highly persistent, with estimated $\hat{\rho}_{MUB}$'s being equal to 0.67, 0.69, 0.85 and 0.9, respectively. For Malaysia, the Netherlands and South Africa, the short sample periods, together with the comparatively high persistence of the CCRs, suggest that results from Johansen's tests are not incompatible with the presence of cointegration. In these three cases it is therefore not clear whether statistical tests have captured the truth, or whether there might indeed be cointegration in the data. A similar argument can be made for the United States. Although the sample period is comparatively longer, the relatively high persistence of the CCR, with $\hat{\rho}_{MUB} = 0.9$, makes it hard for Johansen's tests to detect cointegration. The overall evidence for the United States is thus not entirely incompatible with the presence of cointegration.

Again, results based on the Selden-Latané specification are very similar to those produced by the semi-log one. In particular, for Finland (1914-1985), Malaysia, the Netherlands, Portugal, Sweden and the United States both Johansen's and Shin's tests indicate absence of cointegration. For South Africa, on the other hand, Johansen's tests produce contrasting evidence. Together with the stationary-looking

[37] These results refer to the Friedman-Schwartz M_2 aggregate. For the Lucas-Nicolini aggregate the results are different in that Shin's test cannot reject cointegration. Given the previously discussed lower extent of informativeness of Shin's tests, these result should however be discounted.

CCR in figure B.6, this provides some weak evidence in favor of cointegration. For Jordan, based on the Selden-Latané specification, Shin's test also indicates no cointegration, whereas Johansen's tests are in favor of cointegration. The clearly non-stationary-looking cointegration residual in figure B.6, however, casts doubt on the whether cointegration truly is in the data.

Finally, based on the log-log specification, Johansen's and Shin's tests both point towards no cointegration for Denmark and Sweden under the Gold Standard, and for Canada, Chile, Jordan, Lebanon and Malaysia for the period since 1914. Indeed, for Canada, Chile and Jordan the CCRs, shown in figures B.3 and B.4, do not look stationary, and they all exhibit high persistence, with the 90 percent confidence interval for $\hat{\rho}_{MUB}$ stretching up to 1.01 in all three cases. Overall, it is thus reasonable to interpret the entire body of evidence as pointing towards absence of cointegration. For Denmark and Sweden under the Gold Standard and for Lebanon evidence is less clear-cut. In those three cases the CCRs look borderline stationary (see Figures B.3 and B.4, respectively). The rather short sample periods and the non-negligible persistence in the CCRs make it hard for Johansen's tests to detect cointegration. In these cases evidence thus is not entirely incompatible with the presence of cointegration.

6.2 Unrestricted specifications

The results produced by unrestricted specifications are reported in tables 2 and 3, whereas the corresponding CCRs are shown in Figures B.7 and B.8. Table 2 reports results from cointegration tests based on unrestricted specifications featuring the logarithms of nominal M_2 and nominal GDP, and either the logarithm of the short rate (for high-inflation countries), or its level (for all other countries). As previously mentioned, in several cases it is not possible to estimate these specifications because for at least one series the order of integration is greater than one. Table 3 reports results from cointegration tests based on unrestricted specifications featuring the logarithms of real M_2 and real GDP, and either the logarithm of the short rate (for high-inflation countries), or its level (for all other countries).

6.2.1 Evidence in favor of cointegration

Based on the nominal specification evidence points towards cointegration for Finland (1946-1985) and Norway (1946-2013). Based on the real specification both Johansen's and Shin's tests provide evidence of cointegration for Korea and for the United States

Table 1: Cointegration tests: Johansen's and Shin's test statistics for all specifications based on M_2 velocity and the short-term interest rate

		Semi-log specification			**Log-log specification**			**Selden-Latané specification**		
		Johansen's tests		Shin's tests	Johansen's tests		Shin's tests	Johansen's tests		Shin's tests
Country	*Period*	*Trace test*	*Max. eig.*		*Trace test*	*Max. eig.*		*Trace test*	*Max. eig.*	
Argentina	1914-1984				12.120 (0.254)	9.773 (0.304)	0.463 (0.297)			
Brazil	1934-2012				8.332 (0.611)	6.693 (0.641)	0.384 (0.330)			
Belize	1977-2014	12.148 (0.294)	11.555 (0.209)	0.266 (0.159)				14.828 (0.157)	13.052 (0.132)	0.226 (0.148)
Canada	1926-2006	12.444 (0.321)	8.293 (0.499)	0.253 (0.547)	15.667 (0.152)	12.672 (0.161)	1.568 (0.029)	11.802 (0.357)	7.650 (0.560)	0.270 (0.534)
Chile	1941-1995	10.167 (0.476)	7.812 (0.522)	0.988 (0.011)	7.978 (0.708)	6.311 (0.726)	0.751 (0.079)			
Colombia	1959-2009	6.534 (0.750)	6.351 (0.670)	0.433 (0.490)	9.761 (0.424)	9.002 (0.383)	0.158 (0.324)	7.621 (0.636)	6.566 (0.645)	0.442 (0.454)
Denmark	1875-1913	15.716 (0.131)	10.797 (0.281)	0.318 (0.268)	14.415 (0.177)	9.860 (0.355)	0.348 (0.043)	15.998 (0.129)	9.822 (0.370)	0.321 (0.217)
	1923-2011	12.358 (0.334)	10.315 (0.312)	0.721 (0.116)	11.310 (0.396)	10.972 (0.260)	0.343 (0.457)	12.273 (0.340)	10.000 (0.344)	0.687 (0.141)
Ecuador	1965-2011	5.349 (0.877)	4.306 (0.897)	0.591 (0.165)	6.878 (0.753)	5.740 (0.761)	0.265 (0.346)	8.440 (0.643)	4.297 (0.917)	0.370 (0.211)
Finland	1914-1985	9.259 (0.553)	5.667 (0.777)	1.369 (0.042)	9.555 (0.536)	5.797 (0.770)	0.461 (0.245)	8.159 (0.661)	4.440 (0.895)	1.448 (0.041)
	1946-1985	13.842 (0.253)	13.681 (0.146)	0.124 (0.188)				21.812 (0.035)	19.162 (0.032)	0.129 (0.225)
Germany	1876-1913	14.703 (0.147)	13.377 (0.136)	0.639 (0.204)	13.836 (0.175)	12.411 (0.172)	0.121 (0.124)	16.202 (0.156)	11.890 (0.270)	0.640 (0.159)
	1960-1989	17.206 (0.145)	16.890 (0.077)	0.425 (0.184)	17.670 (0.130)	17.370 (0.067)	0.049 (0.709)	17.761 (0.131)	17.133 (0.074)	0.400 (0.217)
Guatemala	1980-2012	13.906 (0.330)	11.142 (0.355)	0.148 (0.284)	13.953 (0.248)	10.228 (0.347)	0.088 (0.678)	14.085 (0.308)	11.340 (0.335)	0.148 (0.298)
Israel	1983-2013	166.062 (0.000)	163.862 (0.000)	0.134 (0.404)	11.796 (0.454)	9.573 (0.505)	0.176 (0.104)			
Italy	1861-1913	19.599 (0.046)	14.904 (0.083)	0.249 (0.245)	17.373 (0.085)	13.100 (0.137)	0.573 (0.069)			
Japan	1955-2013	11.603 (0.289)	11.422 (0.215)	0.408 (0.141)	6.195 (0.782)	6.189 (0.695)	0.293 (0.101)			
Jordan	1965-2015	12.258 (0.296)	8.292 (0.502)	1.299 (0.015)	12.357 (0.262)	8.246 (0.466)	0.298 (0.069)	28.714 (0.002)	24.515 (0.004)	1.266 (0.012)
Korea	1960-2014	30.723 (0.002)	27.195 (0.002)	0.332 (0.250)	12.814 (0.229)	10.089 (0.303)	0.170 (0.330)	60.709 (0.000)	45.390 (0.001)	0.093 (0.820)
Lebanon	1977-2014	17.831 (0.156)	14.222 (0.173)	0.151 (0.262)	16.827 (0.196)	13.124 (0.231)	0.921 (0.008)	19.479 (0.083)	15.650 (0.096)	0.141 (0.229)
Malaysia	1980-2012	9.843 (0.448)	9.093 (0.406)	0.556 (0.070)	9.668 (0.476)	8.911 (0.434)	0.164 (0.096)	11.090 (0.390)	8.596 (0.474)	0.560 (0.065)
Mexico	1985-2014	27.628 (0.009)	27.553 (0.006)	0.185 (0.314)	11.024 (0.447)	11.021 (0.337)	0.111 (0.193)	28.108 (0.009)	27.623 (0.006)	0.116 (0.640)
Netherlands	1957-1992	11.522 (0.474)	10.556 (0.395)	0.619 (0.001)	13.317 (0.339)	12.297 (0.264)	0.101 (0.275)	11.311 (0.486)	10.790 (0.369)	0.613 (0.004)
Peru	1960-2014				9.386 (0.532)	8.701 (0.431)	0.464 (0.115)			
Paraguay	1962-2014	21.367 (0.030)	17.475 (0.038)	0.259 (0.426)				21.287 (0.035)	16.187 (0.062)	0.241 (0.425)
Portugal	1914-1998	8.137 (0.653)	4.219 (0.916)	2.525 (0.011)	6.431 (0.828)	4.086 (0.933)	1.033 (0.174)	8.558 (0.636)	4.915 (0.869)	2.471 (0.005)
South Africa	1967-2014	14.387 (0.239)	13.118 (0.180)	0.641 (0.043)	15.642 (0.141)	15.421 (0.070)	0.234 (0.269)	16.620 (0.131)	15.364 (0.085)	0.631 (0.022)
Spain	1946-1997	9.391 (0.568)	7.041 (0.628)	0.422 (0.231)	8.419 (0.665)	6.175 (0.738)	0.617 (0.082)	9.177 (0.586)	6.928 (0.642)	0.420 (0.230)
Sweden	1871-1913	11.131 (0.395)	10.443 (0.344)	0.691 (0.210)	11.168 (0.395)	10.462 (0.344)	0.171 (0.088)	10.755 (0.420)	8.474 (0.447)	2.203 (0.010)
	1914-2010	9.921 (0.500)	7.549 (0.555)	2.189 (0.011)	4.884 (0.919)	4.797 (0.851)	0.333 (0.560)			
Taiwan	1962-2014	20.774 (0.021)	15.719 (0.056)	0.357 (0.164)	13.506 (0.169)	8.617 (0.402)	0.215 (0.140)			
United States	1869-1913	12.673 (0.218)	12.480 (0.147)	0.173 (0.915)	13.973 (0.170)	13.697 (0.109)	0.044 (0.874)	14.669 (0.151)	13.605 (0.117)	0.161 (0.887)
FS M_2	1915-2014	12.794 (0.280)	8.195 (0.485)	1.636 (0.022)	7.247 (0.733)	5.344 (0.801)	0.392 (0.538)	11.274 (0.387)	7.244 (0.589)	1.765 (0.024)
LN M_2	1915-2014	11.794 (0.333)	7.582 (0.548)	0.408 (0.426)	7.559 (0.699)	6.989 (0.612)	0.549 (0.290)	13.570 (0.217)	8.184 (0.472)	0.432 (0.397)

Note: Bootstrapped p-values (in parentheses) are based on 10,000 bootstrap replications.

under the Gold Standard. Based on the nominal specification,[38] however, Johansen's tests produce contrasting results for the United States under the Gold Standard, with the maximum eigenvalue test pointing towards cointegration, and the trace test being instead marginally insignificant (with a p-value of 0.105). Shin's test also points towards no cointegration. The CCR, however, appears stationary and thus suggests that a cointegration relationship still might be there.

6.2.2 Evidence against cointegration

Based on the nominal specification, for Denmark (1923-2011) evidence from statistical tests clearly speaks against cointegration, and the strong look of non-stationarity of the CCRs only confirms this. Based on the real specification Shin's test is not able to reject the null of cointegration for Denmark, however, the CCRs look non-stationary as in the nominal specification which confirms the result based on the nominal specification. For Canada and Chile (1915-1995), for which cointegration could only be tested based on the real specification, all tests point towards no cointegration, a result which is validated by the look of non-stationarity of the CCRs.

6.2.3 Conflicting evidence

Based on the nominal specification, for Ecuador, Finland (1914-1985), Portugal (1914-1998), Sweden (1914-2010) and the United States (1915-2014), and Italy and Sweden under the Gold Standard, Johansen's and Shin's test produce contrasting evidence. For Finland, the United States, and Italy and Sweden under the Gold Standard, CCRs appear as stationary, although in the three former cases they are very highly persistent. Thus, cointegration might be in the data but Johansen's tests might not be able to detect it. The high estimates of $\hat{\rho}_{MUB}$ for Finland and the United States for the period since 1914, and the comparatively short sample period for Italy and Sweden under the Gold Standard are also compatible with this position. For Ecuador, Portugal and Sweden (1914-2010) it appears as less likely that a cointegration relationship exists, since the CCRs appear as non-stationary. Mixed evidence also results for South Africa, and for the United States under the Gold Standard. Although for South Africa both Johansen's trace test and Shin's test point towards cointegration, the null of no cointegration cannot be rejected based on the maximum eigenvalue test. Further, the rather stationary-looking CCR suggests

[38]For Korea cointegration tests could not be performed based on the unrestricted nominal specification.

that cointegration might be there.

Based on the real specification, for Israel the maximum eigenvalue test and Shin's test point towards cointegration, whereas the trace test is instead marginally insignificant (with a p-value of 0.107). For Denmark under the Gold Standard, Shin's test and the trace test provide evidence of cointegration, whereas the maximum eigenvalue test cannot reject the null of no cointegration.

Table 2: Cointegration tests: Johansen's and Shin's test statistics for the logarithm of nominal M_2, nominal GDP and the short-term interest rate

		I: Johansen's tests			II: Shin's tests
		Trace tests	*Maximum eigenvalue tests*		
Country	*Period*		0 *versus* 1	1 *versus* 2	
Belize	1977-2014	39.270 (0.040)	27.794 (0.043)	7.829 (0.532)	0.148 (0.057)
Denmark	1923-2011	22.341 (0.547)	14.693 (0.529)	5.442 (0.832)	0.249 (0.068)
Ecuador	1965-2011	10.974 (0.991)	6.951 (0.986)	3.879 (0.934)	0.087 (0.355)
Finland	1914-1985	16.588 (0.871)	10.994 (0.827)	4.987 (0.829)	0.031 (0.998)
	1946-1985	44.643 (0.023)	29.343 (0.042)	14.884 (0.138)	0.043 (0.663)
Italy	1861-1913	24.653 (0.355)	16.878 (0.322)	6.868 (0.578)	0.098 (0.488)
Norway	1946-2013	42.381 (0.008)	29.500 (0.013)	11.224 (0.176)	0.077 (0.638)
Portugal	1914-1998	18.728 (0.799)	9.795 (0.921)	7.402 (0.613)	0.156 (0.385)
South Africa	1967-2014	41.060 (0.044)	22.833 (0.167)	9.900 (0.425)	0.040 (0.847)
Sweden	1871-1913	28.593 (0.338)	20.025 (0.280)	8.490 (0.501)	0.097 (0.419)
	1914-2010	14.170 (0.915)	9.026 (0.903)	4.259 (0.874)	0.178 (0.455)
United States	1869-1913	32.439 (0.105)	22.816 (0.092)	9.570 (0.354)	0.175 (0.013)
Friedman-Schwartz M_2	1915-2014	13.522 (0.935)	9.331 (0.884)	3.624 (0.938)	0.116 (0.782)
Lucas-Nicolini M_2	1915-2014	16.777 (0.797)	10.218 (0.828)	6.298 (0.667)	0.296 (0.198)

Note: Test statistics are from the regression $\ln(M_{2,t}) = \beta_0 + \beta_1 \ln(NGDP_t) + \beta_2 R_t + u_t$, where $M_{2,t}$ = nominal M_2; $NGDP_t$ = nominal GDP; R_t = short rate. The bootstrapped p-values (in parentheses) are based on 10,000 bootstrap replications of the VECM estimated under the null hypothesis of one cointegration vector.

6.3 Discussion of the evidence

Overall, only in very few cases do statistical tests provide strong evidence of a stable demand for M_2. In particular, Finland (1946-1985), Korea, Mexico and Paraguay are the only countries for which all tests consistently point towards cointegration based on more than one money demand specification. Specifically, for Finland this is the case based on the Selden-Latané and the unrestricted nominal specification. Based on the semi-log specification the results of the cointegration tests are mixed. For Korea cointegration tests consistently point towards cointegration based on the semi-log, the Selden-Latané, and the unrestricted real specification, whereas the only specification producing mixed evidence is the log-log. As discussed above, the log-log specification does not appear to work well for any country in the dataset. For Mexico and Paraguay a cointegration relation is found based on the semi-log

Table 3: Cointegration tests: Johansen's and Shin's test statistics for the logarithm of real M_2, real GDP and the short-term interest rate

		I: Johansen's tests			II: Shin's tests
		Trace tests	*Maximum eigenvalue tests*		
Country	*Period*		0 *versus* 1	1 *versus* 2	
Canada	1926-2006	15.445 (0.885)	9.490 (0.895)	5.930 (0.725)	0.257 (0.180)
Chile*	1915-1995	20.676 (0.541)	13.943 (0.501)	5.661 (0.759)	0.402 (0.015)
Colombia	1959-2009	25.653 (0.322)	16.175 (0.385)	7.402 (0.627)	0.043 (0.776)
Denmark	1875-1913	34.903 (0.079)	21.429 (0.155)	13.467 (0.105)	0.053 (0.696)
	1923-2011	16.968 (0.797)	11.596 (0.725)	3.620 (0.938)	0.177 (0.282)
Finland	1946-1985	45.656 (0.007)	28.323 (0.028)	13.596 (0.117)	0.093 (0.162)
Israel*	1983-2013	42.045 (0.107)	36.348 (0.025)	5.165 (0.898)	0.027 (0.959)
Italy	1861-1913	37.153 (0.054)	22.931 (0.114)	14.155 (0.091)	0.178 (0.092)
Korea	1960-2014	46.983 (0.009)	22.851 (0.100)	13.812 (0.102)	0.079 (0.746)
Portugal	1914-1998	24.767 (0.345)	14.481 (0.498)	8.129 (0.441)	0.196 (0.293)
South Africa	1967-2014	33.714 (0.145)	21.000 (0.222)	12.681 (0.236)	0.091 (0.230)
Spain	1946-1997	14.863 (0.914)	10.153 (0.859)	3.920 (0.930)	0.071 (0.606)
Sweden	1871-1913	34.313 (0.122)	23.069 (0.142)	9.617 (0.398)	0.061 (0.536)
Sweden	1914-2010	12.694 (0.950)	9.420 (0.859)	3.236 (0.954)	0.246 (0.225)
United States	1869-1913	48.387 (0.006)	32.365 (0.014)	12.405 (0.209)	0.054 (0.713)
Friedman-Schwartz M_2	1915-2013	15.262 (0.874)	10.380 (0.821)	4.019 (0.897)	0.176 (0.424)
Lucas-Nicolini M_2	1915-2013	19.326 (0.645)	10.871 (0.785)	7.642 (0.465)	0.330 (0.107)
Venezuela*	1962-1998	19.397 (0.771)	11.594 (0.817)	7.598 (0.589)	0.043 (0.745)

Note: Test statistics are from the regression $\ln(M_{2,t}) = \beta_0 + \beta_1 \ln(RGDP_t) + \beta_2 R_t + u_t$, where $M_{2,t}$ = real M_2; $RGDP_t$ = real GDP; R_t = short rate. The bootstrapped p-values (in parentheses) are based on 10,000 bootstrap replications of the VECM estimated under the null hypothesis of one cointegration vector.
* Short rate is in logarithm.

and the Selden-Latané specification. Based on the log-log specification the evidence is mixed for Mexico. For Paraguay cointegration tests could not be conducted based on the log-log specification. For Taiwan, cointegration tests could only be conducted based on the semi-log and the log-log specification, and only based on the semi-log do they produce evidence of cointegration. For Norway cointegration could be tested based on only one specification, which is the unrestricted nominal specification. Based on this specification a cointegrating relation is detected. For Israel and Lebanon, and for Italy and the United States under the Gold Standard, at least one money demand specification produces evidence of cointegration. For Israel, cointegration is detected based on the semi-log specification, whereas weak evidence of cointegration is produced by the trivariate real specification. Mixed evidence exists based on the log-log specification. For the other two countries, evidence is mixed across the alternative specifications. For example, for Lebanon evidence is in favor of cointegration based on the Selden-Latané specification, whereas the semi-log specification produces mixed results and the log-log specification produces evidence against cointegration.

On the other hand, evidence against cointegration is strong for Canada, Chile (1941-

1995), Finland (1914-1985), Jordan, Portugal, and Sweden. For all these cases both Johansen's and Shin's tests point towards no cointegration based on more than one specification, and the CCRs clearly appear as non-stationary. For Malaysia, the Netherlands and the United States (1915-2014), the tests also uniformly point towards no cointegration based on the semi-log and the Selden-Latané specification. For Malaysia, also the log-log specification points towards no cointegration. However, the cointegration residuals do rather seem to be stationary. Therefore, we cannot be entirely sure that there truly is no cointegration relation in the data.

6.4 Possible explanations for the instability of M_2 demand

The evidence discussed above shows that it is almost impossible to detect a stable demand for M_2. Although M_2 may better internalize portfolio shifts between M_1 and the non-M_1 M_2 components, for a majority of countries in the dataset, long-run demand for M_2 is unstable. For M_1, on the other hand, Benati et al. (2018) detect a stable long-run demand in many countries. In this section, I discuss possible reasons for the instability in the long-run demand for M_2. The fact that the M_1-part of M_2 exhibits a stable long-run demand implies that the instability of M_2 has to come from the non-M_1 component of M_2.[39] While the components of the narrow M_1 aggregate mostly fulfill the role of a medium of exchange to facilitate transactions, the non-M_1 components of M_2 rather have the role of an asset. Further, the non-M_1 components usually pay interest, whereas the M_1 components, i.e. currency and demand deposits, do not.

In the literature a number of different explanations for the instability in the demand for money have been discussed. One strand of the literature has proposed the so-called institutional approach which has been revisited by Bordo and Jonung (1990, 1997, 2009) and Siklos (1993). This approach emphasizes financial development and institutional changes within the economy as a major influence on money velocity. Bordo and Jonung have observed that money velocity behaves differently across countries depending on their stage of financial and economic development. Stressing the role of M_2 as an asset, they discuss two different processes, besides movements in the interest rate, that are responsible for driving the behavior of velocity. The first process is financial deepening and the second one is financial innovation - especially the creation of money substitutes. In addition, they stress the influence of improved economic stability.

[39] i.e. savings deposits and, for some countries, time deposits.

Financial deepening

Financial deepening is the first stage of the financial development process. With the development of financial institutions and the growth of the financial sector more people get access to financial services. Through the rise of a commercial banking system that supplies notes and deposit facilities to the public the use of money for settling transactions grows. These developments promote a more rapid growth in the demand for money than in nominal income and thus cause money velocity to fall. Bordo and Jonung explain the downward trend in money velocity for the industrialized countries until the interwar period (see figure 3) with this process of monetization. For industrialized countries, such as Canada and the United States or Denmark, Sweden and Norway, financial deepening essentially ended around World War I. For several other countries in the dataset, a downward trend in M_2 velocity can be observed until recently, indicating that the process of financial deepening started later and only ended recently. For example, for developing countries such as Colombia, Ecuador, Jordan, Malaysia, South Africa and Venezuela M_2 velocity keeps decreasing until lately, while the interest rate starts increasing during the 1980s. But also for some developed countries, such as Italy, Portugal and Spain, the process of financial deepening took longer, with velocity exhibiting a downward trend until the 1970s.

Financial innovation

The second stage of financial development is increasing financial sophistication due to financial innovation. Financial innovation is a popular explanation for the postwar rise in money velocity that can be observed for most industrialized countries. Financial sophistication has emerged twofold. On the one hand, a large number of new financial instruments have emerged that act as close substitutes for money, such as stocks, bonds, and other financial assets. These money substitutes may satisfy asset motives previously met by deposits, and thus reduce demand for money as an asset at any given interest rate, shifting the velocity curve upward. On the other hand, various methods of economizing on money balances, such as the use of credit cards or electronic cash management techniques have facilitated transactions. These modern money transfer methods reduce the transactions demand for money and cause a rise in money velocity.

Economic stability

Improved economic security and stability are other factors causing velocity to rise. Economic security provided by a modern welfare state, such as unemployment benefits, public insurance schemes, and pension systems reduce the need of holding money as a store of value for cases of financial straits. Improved economic stability as a consequence of macroeconomic policies aimed at maintaining maximum employment and an environment of price stability further reduces demand for money.

The institutional approach suggests that additional explanatory variables to the standard determinants of velocity used in the money demand literature are needed to explain the behavior of money velocity. Siklos (1993) has tested the institutional hypothesis for five industrialized countries since the end of the 19th century up to the mid-1980s using cointegration methods. He includes a number of different proxy variables in the money demand equation to account for financial innovation, economic stability, and the process of monetization. For all countries, except the United States,[40] he has only found a stable long-run M_2 demand once the institutional variables are added to the money demand equation.
More recently, studies focusing on the United States since WWII have modeled financial innovation through the use of time trends or dummy and shift variables. In the early nineties M_2 growth began to slow down despite a considerable reduction in the interest rate. This change in the relationship between M_2 velocity and interest rates has given rise to different explanations as to the causes of this change. A widely stated cause of the shift in M_2 velocity has been attributed to the development of bond and stock mutual funds (Darin and Hetzel, 1994; Mehra, 1997; Duca, 2000; Duca and VanHoose, 2004). Bond and stock funds have facilitated accessibility to stocks and bonds for the public. They have reduced transaction costs of switching between M_2 and securities not included in M_2, leading to a higher substitutability between M_2 and stock and bond funds.[41] Some papers have modeled financial innovation using shift dummies, for example in the form of a broken linear trend. By adding these ad-hoc explanatory variables, stability of M_2 demand could be restored, indicating that the causes discussed above are reasonable explanations for the recent instability in U.S. broad money demand.

[40] For the United States, he has found a stable demand for M_2 using the traditional money demand equation.

[41] Another related explanation for the instability in U.S. M_2 demand is the poor financial condition of U.S. depository institutions, especially thrift institutions, in the early 1990s (Lown et al., 2006; Carlson et al., 2000), leading to the development of mutual funds.

Other attempts to restore stability in long-run M_2 demand aimed at adjusting M_2. Financial innovation caused money substitutes to become more attractive. This has led many authors to propose a change of the M_2 definition. For example, it was suggested to exclude small-denomination time deposits from M_2 (Carlson et al., 2000; Lown et al., 2006), or to include money substitutes, such as stock and bond mutual funds (Orphanides et al., 1994; Darin and Hetzel, 1994). However, these adjustments have only worked temporarily, and the additional components have been subject to new problems.[42]

7 Conclusion

This paper has reexamined one of the central questions in macroeconomic research over the past decades: is there a stable long-run money demand? I used a novel dataset on 32 countries to explore this question for the M_2 aggregate. I analyzed five money demand specifications, including three restricted specifications which are based on money velocity and a short term nominal interest rate, and two unrestricted specifications which feature nominal or real M_2 and GDP and a short-term nominal interest rate. I tested for the stability of money demand using two different cointegration tests: Johansens' tests and Shin's test. Because the persistence in the CCR plays a crucial role for the performance of the tests, all results have been analyzed taking the estimated persistence into account.

The main result obtained from the analysis is the following. Evidence of cointegration is hard to find in a majority of the cases. Based on the entire evidence, including the persistence of the candidate cointegration residuals, the results clearly speak against a stable money demand for Canada, Chile (1941-1995), Finland (1914-1985), Jordan, Portugal, and Sweden. Only in five cases, Finland (1946-1985), Korea, Mexico, Paraguay and Taiwan, is there strong evidence in favor of cointegration. Weak evidence exists for Israel and Lebanon. For the remaining countries, the evidence is mixed across specifications, or across the two tests and with regard to the evidence from the cointegration residuals.

Further, the evidence suggests that the log-log specification does not work well as a model of the long-run demand for M_2 for any country in the dataset. In fact, it is the only specification under consideration based on which the cointegration tests do not produce consistent evidence in favor of cointegration for neither country.

[42]For example, Orphanides et al. (1994) noticed that adding stock and bond mutual funds makes the new aggregate very sensitive to movements in bond and equity price.

Comparing these results for M_2 to results obtained by Benati et al. (2018) for M_1, it is clear that a stable long-run money demand relation is much harder to find for M_2 than for M_1. Therefore, if any monetary aggregate is to be used as an indicator for nominal economic activity, the M_1 aggregate is likely to be the better choice.
The different results for the two monetary aggregates prompt the question: what causes these differences? Since the M_1 aggregate is included in the M_2 aggregate, the instability in M_2 demand arises from the non-M_1 component of M_2, i.e. M_2-M_1. It is therefore important to split the broader aggregate into its stable part (M_1) and its unstable part (M_2-M_1) to understand the demand for money. This is what I am pursuing in the next chapter of this dissertation.

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A Additional tables

Table A.1: Persistence of the candidate cointegration residuals for the bivariate money demand specifications (median, and 90 percent bootstrapped confidence interval)

		Money demand specification		
Country	*Period*	*Semi-log*	*Log-log*	*Selden-Latané*
Argentina	1914-1984		1.01 [0.99;1.05]	
Belize	1977-2014	0.71 [0.50;0.97]		0.70 [0.50;1.00]
Brazil	1934-2012		0.96 [0.87;1.02]	
Canada	1926-2006	0.92 [0.85;1.01]	0.92 [0.84;1.00]	0.92 [0.85;1.01]
Chile	1941-1995	0.68 [0.53;0.84]	0.80 [0.62;1.02]	
Colombia	1959-2009	0.88 [0.76;1.11]	0.86 [0.68;1.14]	0.87 [0.74;1.01]
Denmark	1923-2011	0.91 [0.85;0.98]	0.88 [0.82;0.96]	0.91 [0.85;0.98]
Ecuador	1965-2011	0.90 [0.75;1.02]	1.00 [0.94;1.03]	0.92 [0.75;1.02]
Finland	1914-1985	0.97 [0.92;1.01]	0.97 [0.93;1.01]	0.98 [0.88;1.02]
Finland	1946-1985	0.54 [0.30;0.83]		0.55 [0.31;0.81]
Germany	1960-1989	0.30 [-0.02;0.62]	0.27 [-0.04;0.59]	0.28 [-0.04;0.62]
Guatemala	1980-2012	0.82 [0.56;1.03]	0.76 [0.50;1.02]	0.83 [0.56;1.03]
Israel	1983-2013	0.36 [0.33;0.40]	0.75 [0.56;1.00]	
Japan	1955-2013	0.63 [0.47;0.79]	0.87 [0.79;0.97]	
Jordan	1965-2015	0.98 [0.93;1.01]	0.81 [0.67;1.01]	0.97 [0.92;1.00]
Korea	1960-2014	0.69 [0.55;0.84]	0.83 [0.72;0.97]	0.72 [0.58;0.87]
Lebanon	1977-2014	0.81 [0.66;1.01]	0.80 [0.64;0.99]	0.57 [0.35;0.80]
Malaysia	1980-2012	0.67 [0.44;0.99]	0.66 [0.40;0.99]	0.69 [0.47;0.99]
Mexico	1985-2014	0.67 [0.50;0.89]	0.58 [0.28;0.94]	0.63 [0.44;0.86]
Netherlands	1957-1992	0.69 [0.46;1.00]	0.75 [0.48;1.02]	0.69 [0.44;1.01]
Paraguay	1962-2014	0.78 [0.61;1.01]		0.78 [0.61;1.01]
Peru	1960-2014		0.78 [0.62;0.96]	
Portugal	1914-1998	0.98 [0.93;1.01]	1.00 [0.96;1.02]	1.00 [0.95;1.02]
South Africa	1967-2014	0.85 [0.71;1.01]	0.86 [0.74;1.01]	0.85 [0.71;1.01]
Spain	1946-1997	0.86 [0.70;1.02]	0.92 [0.76;1.03]	0.85 [0.68;1.02]
Sweden	1914-2010	0.93 [0.86;1.01]	0.94 [0.87;1.01]	0.93 [0.86;1.01]
Taiwan	1962-2014	0.65 [0.48;0.83]	0.82 [0.69;0.97]	
United States	1915-2014	0.90 [0.82;0.99]	0.94 [0.87;1.01]	0.91 [0.83;1.00]
Gold Standard Period				
Denmark	1875-1913	0.76 [0.61;0.93]	0.79 [0.64;0.95]	0.80 [0.65;0.96]
Germany	1876-1913	0.38 [0.06;0.70]	0.42 [0.11;0.75]	0.81 [0.63;1.01]
Italy	1861-1913	0.78 [0.66;0.91]	0.81 [0.69;0.93]	
Sweden	1871-1913	0.59 [0.41;0.78]	0.60 [0.41;0.79]	
United States	1869-1913	0.49 [0.21;0.78]	0.40 [0.11;0.74]	0.48 [0.20;0.78]

Note: Hansen (1999) 'grid bootstrap' estimates of the sum of the autoregressive coefficients based on AR(2) models are based on 2,000 bootstrap replications for each value of ρ in the grid. Candidate cointegration residuals have been computed based on the respective bivariate model and Johansen's estimator.

Table A.2: Persistence of the candidate cointegration residuals for the trivariate money demand specifications (median, and 90 percent bootstrapped confidence interval)

		Money demand specification	
Country	*Period*	*nominal*	*real*
Belize	1977-2014	0.34 [0.13;0.55]	
Canada	1926-2006		0.92 [0.85;1.00]
Chile*	1941-1995		0.70 [0.60;0.80]
Colombia	1959-2009		0.95 [0.83;1.02]
Denmark	1923-2011	0.95 [0.90;1.01]	0.91 [0.85;0.98]
Ecuador	1965-2011	0.90 [0.77;1.02]	
Finland	1914-1985	0.81 [0.69;0.94]	
Finland	1946-1985	0.29 [0.01;0.58]	0.37 [0.13;0.63]
Israel	1983-2013		0.36 [0.33;0.40]
Korea	1960-2014		0.66 [0.52;0.79]
Norway	1946-2013	0.86 [0.79;0.94]	
Portugal	1914-1998	0.99 [0.89;1.02]	0.91 [0.81;1.01]
South Africa	1967-2014	0.65 [0.48;0.83]	0.80 [0.66;0.99]
Spain	1946-1997		0.80 [0.63;1.01]
Sweden	1914-2010	0.91 [0.84;0.99]	0.91 [0.85;0.99]
United States	1915-2014	0.86 [0.78;0.95]	0.87 [0.79;0.96]
Venezuela	1962-1998		0.61 [0.34;0.92]
Gold Standard period			
Denmark	1875-1913		0.42 [0.21;0.64]
Italy	1861-1913	0.78 [0.66;0.92]	0.69 [0.54;0.84]
Sweden	1871-1913	0.44 [0.24;0.65]	0.58 [0.33;0.82]
United States	1869-1913	0.31 [0.02;0.63]	0.10 [-0.22;0.40]

Note: Hansen (1999) 'grid bootstrap' estimates of the sum of the autoregressive coefficients based on AR(2) models are based on 2,000 bootstrap replications for each value of ρ in the grid. Candidate cointegration residuals have been computed based on the respective trivariate model and Johansen's estimator.
*Short rate in logarithms.

Table A.3: Bootstrapped p-values for Elliott, Rothenberg, and Stock unit root tests

		Logarithm of:															
		nominal GDP		nominal M_2		real GDP		real M_2		M_2 velocity		short rate		M_2 velocity		short rate	
Country	*Period*	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2
Argentina	1914-1984	0.851	0.999	0.990	0.701					0.791	0.792	1.000	0.999	0.775	0.780	0.999	0.999
Belize	1977-2014	0.238	0.510	0.677	0.723					0.764	0.732	0.948	0.990	0.670	0.643	0.857	0.847
Brazil	1934-2012	0.295	0.592	0.298	0.674					0.713	0.725	0.375	0.341	0.600	0.692	0.007	0.059
Canada	1926-2006	0.117	0.134	0.161	0.218	0.680	0.361	0.683	0.549	0.377	0.192	0.614	0.719	0.383	0.219	0.377	0.478
Chile	1941-1995	0.427	0.466	0.620	0.447					0.067	0.336	0.348	0.252	0.004	0.011	0.136	0.098
	1915-1995					0.177	0.059	0.208	0.582			0.370	0.260			0.086	0.037
Colombia	1959-2009	0.970	0.950	0.954	0.961	0.392	0.627	0.831	0.697	0.757	0.641	0.457	0.496	0.602	0.456	0.435	0.280
Denmark	1923-2011	0.218	0.403	0.018	0.410	0.944	0.965	0.051	0.444	0.172	0.430	0.855	0.814	0.262	0.418	0.693	0.686
Ecuador	1965-2011	0.674	0.673	0.535	0.521					0.611	0.646	0.727	0.826	0.295	0.380	0.336	0.664
Finland	1914-1985	0.289	0.103	0.282	0.122	0.015	0.079	0.214	0.192	0.266	0.211	0.538	0.522	0.392	0.400	0.517	0.515
	1946-1985	0.056	0.599	0.889	0.925	0.592	0.638	0.040	0.346	0.005	0.106	0.055	0.041	0.010	0.107	0.133	0.153
Germany	1960-1989	0.842	0.960	0.759	0.928					0.671	0.865	0.068	0.135	0.601	0.816	0.070	0.129
Guatemala	1980-2012	0.959	0.971	0.982	0.993					0.378	0.498	0.641	0.578	0.380	0.493	0.610	0.591
Israel	1983-2013	0.001	0.161	0.002	0.006	0.871	0.854	0.257	0.005	0.079	0.000	0.183	0.011	0.063	0.000	0.120	0.000
Italy	1948-1998	0.780	0.891	0.972	0.976	0.096	0.636	0.099	0.954	0.141	0.665	0.860	0.901	0.019	0.548	0.808	0.839
Japan	1955-2013	0.212	0.720	0.171	0.772	0.055	0.330	0.014	0.261	0.289	0.448	0.728	0.772	0.012	0.044	0.598	0.580
Jordan	1965-2015	0.735	0.538	0.365	0.717					0.023	0.105	0.294	0.171	0.006	0.019	0.413	0.231
Korea	1960-2014	0.273	0.706	0.353	0.877	0.967	0.977	0.832	0.899	0.275	0.276	0.900	0.903	0.087	0.176	0.621	0.525
Lebanon	1977-2014	0.857	0.898	0.905	0.965					0.332	0.295	0.530	0.486	0.200	0.347	0.587	0.648
Malaysia	1980-2012	0.458	0.533	0.710	0.358					0.263	0.368	0.376	0.384	0.070	0.162	0.340	0.318
Mexico	1985-2014	0.014	0.019	0.049	0.005					0.969	0.666	0.628	0.286	0.918	0.491	0.344	0.026
Morocco	1985-2008	0.301	0.255	0.349	0.634					0.731	0.476	0.863	0.739	0.103	0.021	0.893	0.766
Netherlands	1957-1992	0.979	0.942	0.927	0.847	0.769	0.650	0.803	0.759	0.972	0.982	0.271	0.195	0.961	0.958	0.260	0.169
Norway	1946-2013	0.972	0.989	0.584	0.666	0.106	0.992	0.099	0.048	0.009	0.027	0.503	0.550	0.039	0.079	0.597	0.612
Paraguay	1962-2014	0.820	0.889	0.915	0.780					0.248	0.272	0.028	0.057	0.095	0.184	0.114	0.239
Peru	1960-2014	0.699	0.840	0.697	0.774					0.584	0.603	0.505	0.573	0.396	0.457	0.013	0.020
Portugal	1914-1998	0.635	0.610	0.422	0.330	0.236	0.241	0.300	0.221	0.572	0.473	0.725	0.712	0.594	0.455	0.596	0.468
South Africa	1967-2014	0.985	0.988	0.985	0.982	0.480	0.712	0.508	0.424	0.774	0.873	0.381	0.464	0.755	0.873	0.309	0.331
Spain	1946-1997	0.990	0.980	0.998	0.997	0.984	0.971	0.987	0.978	0.655	0.623	0.639	0.746	0.646	0.643	0.399	0.594
Sweden	1914-2010	0.716	0.742	0.605	0.530	0.914	0.937	0.515	0.556	0.330	0.503	0.745	0.811	0.417	0.546	0.475	0.627
Taiwan	1962-2014	0.438	0.798	0.470	0.852	0.198	0.765	0.562	0.881	0.052	0.107	0.598	0.660	0.001	0.003	0.409	0.507
Turkey	1977-2014	0.958	0.828	0.993	0.906					0.965	0.925	0.577	0.614	0.911	0.864	0.684	0.703
United States	1915-2014	0.679	0.353	0.544	0.233	0.487	0.210	0.399	0.316	0.367	0.296	0.627	0.482	0.413	0.354	0.316	0.317
Venezuela	1962-1998	0.668	0.785	0.434	0.671	0.174	0.399	0.888	0.913	0.569	0.563	0.820	0.786	0.392	0.372	0.786	0.785
Gold Standard period																	
Canada	1873-1907	0.501	0.650	0.265	0.228	0.601	0.524	0.314	0.067	0.420	0.601	0.006	0.157	0.221	0.477	0.005	0.131
Denmark	1875-1913	0.129	0.205	0.097	0.027	0.147	0.116	0.762	0.562	0.167	0.122	0.416	0.268	0.134	0.080	0.374	0.264
Finland	1867-1913	0.788	0.828	0.477	0.575	0.228	0.509	0.323	0.431	0.286	0.275	0.096	0.043	0.036	0.039	0.079	0.048
Germany	1876-1913	0.111	0.891	0.038	0.292					0.509	0.132	0.126	0.233	0.179	0.022	0.145	0.255
Italy	1861-1913	0.954	0.994	0.283	0.699	0.816	0.946	0.093	0.712	0.024	0.045	0.761	0.797	0.009	0.021	0.756	0.788
Norway	1830-1913	0.467	0.383	0.413	0.355	0.218	0.257	0.020	0.045	0.602	0.441	0.014	0.005	0.155	0.039	0.011	0.006
Spain	1874-1913	0.957	0.949	0.385	0.252	0.310	0.456	0.526	0.607	0.618	0.618	0.055	0.021	0.655	0.655	0.057	0.019
Sweden	1871-1913	0.930	0.918	0.337	0.749	0.950	0.911	0.065	0.565	0.033	0.258	0.118	0.028	0.001	0.084	0.112	0.027
United States	1869-1913	0.271	0.403	0.403	0.333	0.598	0.518	0.248	0.432	0.302	0.460	0.038	0.234	0.125	0.250	0.018	0.214

Note: Based on 10,000 bootstrap replications of estimated ARIMA processes. Tests are with an intercept and a time trend for the logarithms of nominal and real GDP and nominal and real M_2, and with an intercept and no time trend for the other series.

Table A.4: Bootstrapped p-values for Elliott, Rothenberg, and Stock unit root tests for the variables in log- and first-differences

		Log-difference of:												*First-difference of:*			
		nominal GDP		nominal M_2		real GDP		real M_2		M_2 velocity		short rate		M_2 velocity		short rate	
Country	*Period*	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2	p=1	p=2
Argentina	1914-1984	0.872	0.920	0.934	0.944					0.000	0.000	0.001	0.002	0.000	0.003	0.038	0.475
Brazil	1934-2012	0.136	0.206	0.128	0.209					0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Belize	1977-2014	0.007	0.037	0.021	0.029					0.003	0.007	0.726	0.725	0.004	0.004	0.006	0.009
Canada	1926-2006	0.004	0.002	0.051	0.130	0.001	0.000	0.000	0.003	0.000	0.000	0.000	0.001	0.000	0.000	0.000	0.000
	1941-1995	0.171	0.079	0.006	0.056					0.000	0.000	0.002	0.001	0.000	0.001	0.006	0.004
Chile	1915-1995					0.000	0.000	0.000	0.000			0.000	0.000			0.001	0.000
Colombia	1959-2009	0.638	0.855	0.236	0.615	0.037	0.056	0.005	0.045	0.002	0.035	0.001	0.007	0.000	0.018	0.009	0.014
Denmark	1923-2011	0.005	0.034	0.012	0.009	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.001
Ecuador	1965-2011	0.006	0.035	0.002	0.009					0.003	0.006	0.001	0.017	0.003	0.004	0.001	0.006
Finland	1914-1985	0.013	0.048	0.010	0.009	0.000	0.002	0.001	0.004	0.000	0.002	0.000	0.003	0.000	0.004	0.000	0.002
	1946-1985	0.005	0.033	0.010	0.032	0.001	0.017	0.002	0.010	0.002	0.013	0.002	0.000	0.002	0.014	0.001	0.002
Germany	1960-1989	0.098	0.247	0.047	0.073					0.008	0.008	0.006	0.078	0.009	0.007	0.004	0.051
Guatemala	1980-2012	0.076	0.133	0.017	0.105					0.004	0.027	0.002	0.057	0.004	0.028	0.002	0.098
Israel	1983-2013	0.013	0.000	0.048	0.000	0.015	0.044	0.007	0.010	0.006	0.007	0.004	0.023	0.004	0.000	0.016	0.053
Italy	1948-1998	0.203	0.569	0.321	0.712	0.097	0.195	0.018	0.451	0.002	0.152	0.002	0.127	0.000	0.254	0.001	0.031
Japan	1955-2013	0.510	0.737	0.513	0.595	0.181	0.515	0.037	0.201	0.000	0.003	0.000	0.004	0.001	0.011	0.000	0.000
Jordan	1965-2015	0.005	0.057	0.211	0.175					0.000	0.009	0.001	0.002	0.000	0.057	0.001	0.002
Korea	1960-2014	0.425	0.663	0.190	0.158	0.015	0.099	0.007	0.008	0.001	0.001	0.000	0.001	0.002	0.000	0.001	0.004
Lebanon	1977-2014	0.117	0.141	0.117	0.390					0.005	0.006	0.007	0.053	0.000	0.006	0.006	0.052
Malaysia	1980-2012	0.005	0.034	0.096	0.141					0.005	0.011	0.005	0.011	0.003	0.006	0.007	0.012
Mexico	1985-2014	0.245	0.002	0.271	0.000					0.006	0.068	0.008	0.038	0.004	0.059	0.006	0.008
Morocco	1985-2008	0.018	0.386	0.155	0.302					0.032	0.362	0.164	0.436	0.016	0.144	0.103	0.314
Netherlands	1957-1992	0.164	0.534	0.016	0.126	0.021	0.228	0.006	0.066	0.003	0.088	0.003	0.002	0.004	0.071	0.002	0.003
Norway	1946-2013	0.000	0.019	0.005	0.052	0.001	0.050	0.000	0.000	0.000	0.005	0.000	0.000	0.000	0.002	0.000	0.001
Paraguay	1962-2014	0.090	0.231	0.033	0.088					0.019	0.035	0.000	0.000	0.019	0.019	0.000	0.000
Peru	1960-2014	0.140	0.082	0.087	0.068					0.001	0.004	0.000	0.002	0.001	0.002	0.001	0.002
Portugal	1914-1998	0.028	0.042	0.017	0.030	0.000	0.000	0.000	0.005	0.001	0.000	0.003	0.084	0.000	0.000	0.001	0.004
South Africa	1967-2014	0.025	0.078	0.007	0.025	0.003	0.015	0.002	0.008	0.001	0.004	0.000	0.002	0.001	0.003	0.000	0.001
Spain	1946-1997	0.007	0.059	0.281	0.568	0.015	0.032	0.015	0.076	0.001	0.010	0.001	0.015	0.001	0.005	0.000	0.003
Sweden	1914-2010	0.001	0.000	0.002	0.002	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.009	0.000	0.000	0.000	0.000
Taiwan	1962-2014	0.200	0.525	0.471	0.518	0.036	0.182	0.026	0.033	0.003	0.006	0.000	0.001	0.012	0.001	0.000	0.000
Turkey	1977-2014	0.683	0.763	0.431	0.658					0.030	0.043	0.021	0.153	0.015	0.017	0.007	0.107
United States	1915-2014	0.000	0.000	0.002	0.002	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Venezuela	1962-1998	0.214	0.422	0.118	0.367	0.032	0.031	0.012	0.086	0.009	0.123	0.064	0.063	0.010	0.129	0.082	0.070
Gold Standard period																	
Canada	1873-1907	0.157	0.085	0.051	0.027	0.055	0.133	0.041	0.132	0.049	0.183	0.000	0.027	0.039	0.149	0.001	0.058
Denmark	1875-1913	0.028	0.012	0.009	0.001	0.002	0.000	0.015	0.019	0.029	0.048	0.006	0.004	0.037	0.063	0.003	0.004
Finland	1867-1913	0.003	0.017	0.004	0.028	0.001	0.010	0.003	0.018	0.001	0.019	0.006	0.024	0.001	0.027	0.004	0.030
Germany	1876-1913	0.126	0.147	0.010	0.023					0.006	0.028	0.001	0.013	0.012	0.086	0.001	0.018
Italy	1861-1913	0.000	0.001	0.001	0.005	0.000	0.002	0.000	0.001	0.000	0.003	0.001	0.001	0.000	0.003	0.000	0.002
Norway	1830-1913	0.000	0.000	0.001	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Spain	1874-1913	0.010	0.095	0.008	0.031	0.000	0.015	0.002	0.022	0.009	0.035	0.002	0.004	0.011	0.038	0.004	0.004
Sweden	1871-1913	0.006	0.010	0.002	0.003	0.002	0.017	0.000	0.005	0.011	0.020	0.003	0.004	0.003	0.000	0.002	0.004
United States	1869-1913	0.002	0.006	0.008	0.010	0.002	0.006	0.004	0.005	0.000	0.001	0.000	0.007	0.001	0.000	0.000	0.005

Note: Based on 10,000 bootstrap replications of estimated ARIMA processes. Tests are with an intercept and a time trend for the logarithms of nominal and real GDP and nominal and real M_2, and with an intercept and no time trend for the other series.

B Additional figures

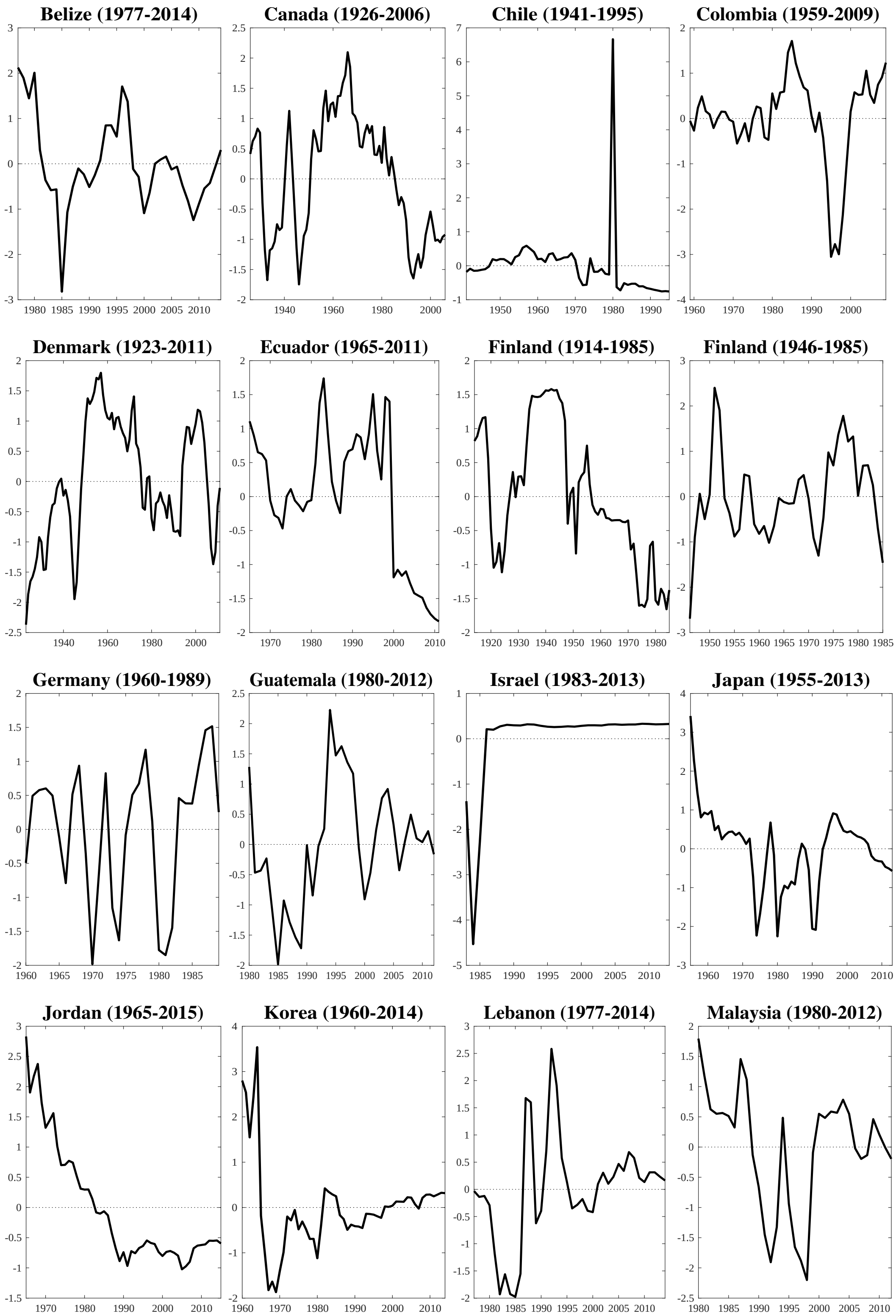

Figure B.1: Cointegration residuals based on the semi-log specification

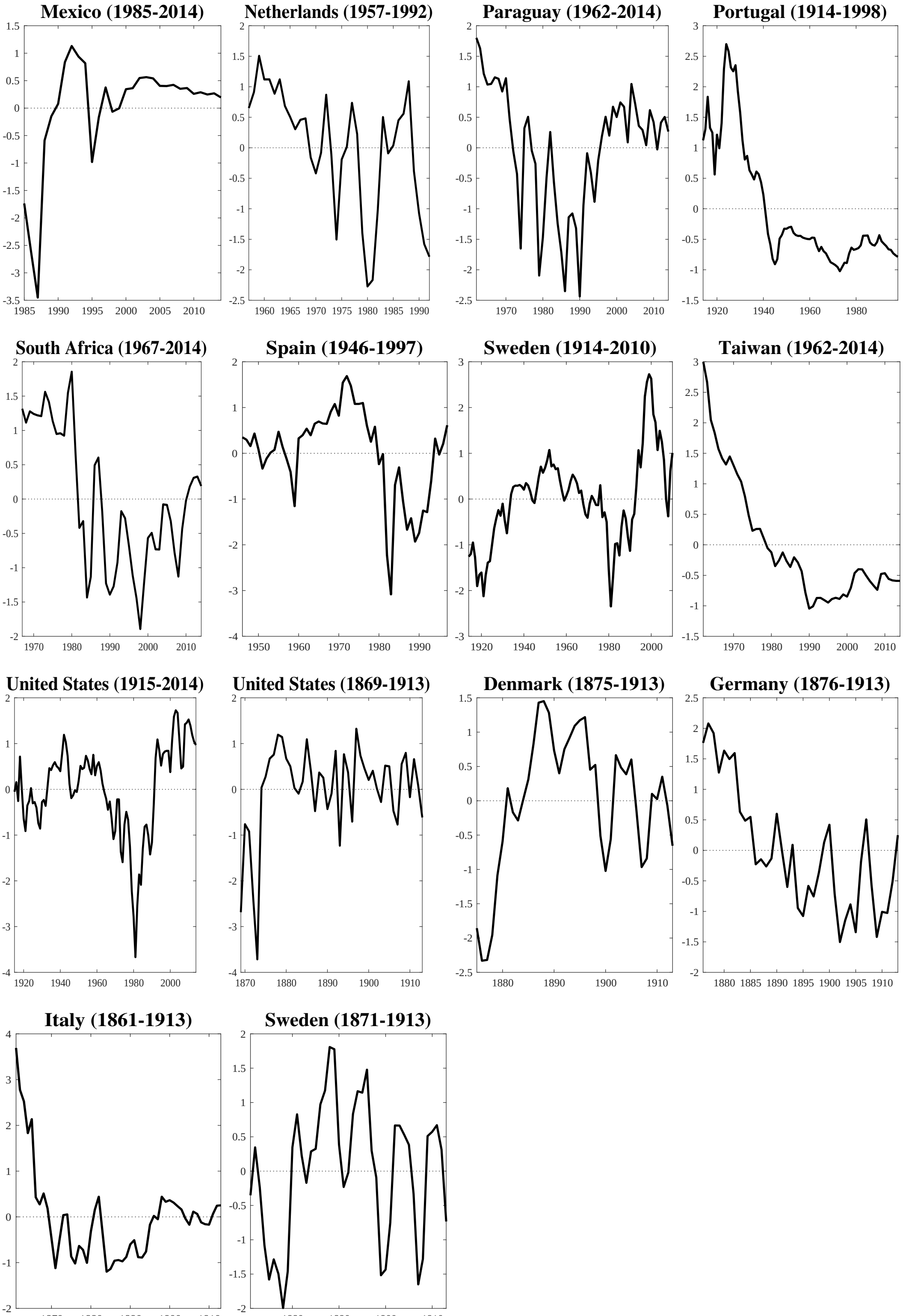

Figure B.2: Cointegration residuals based on the semi-log specification

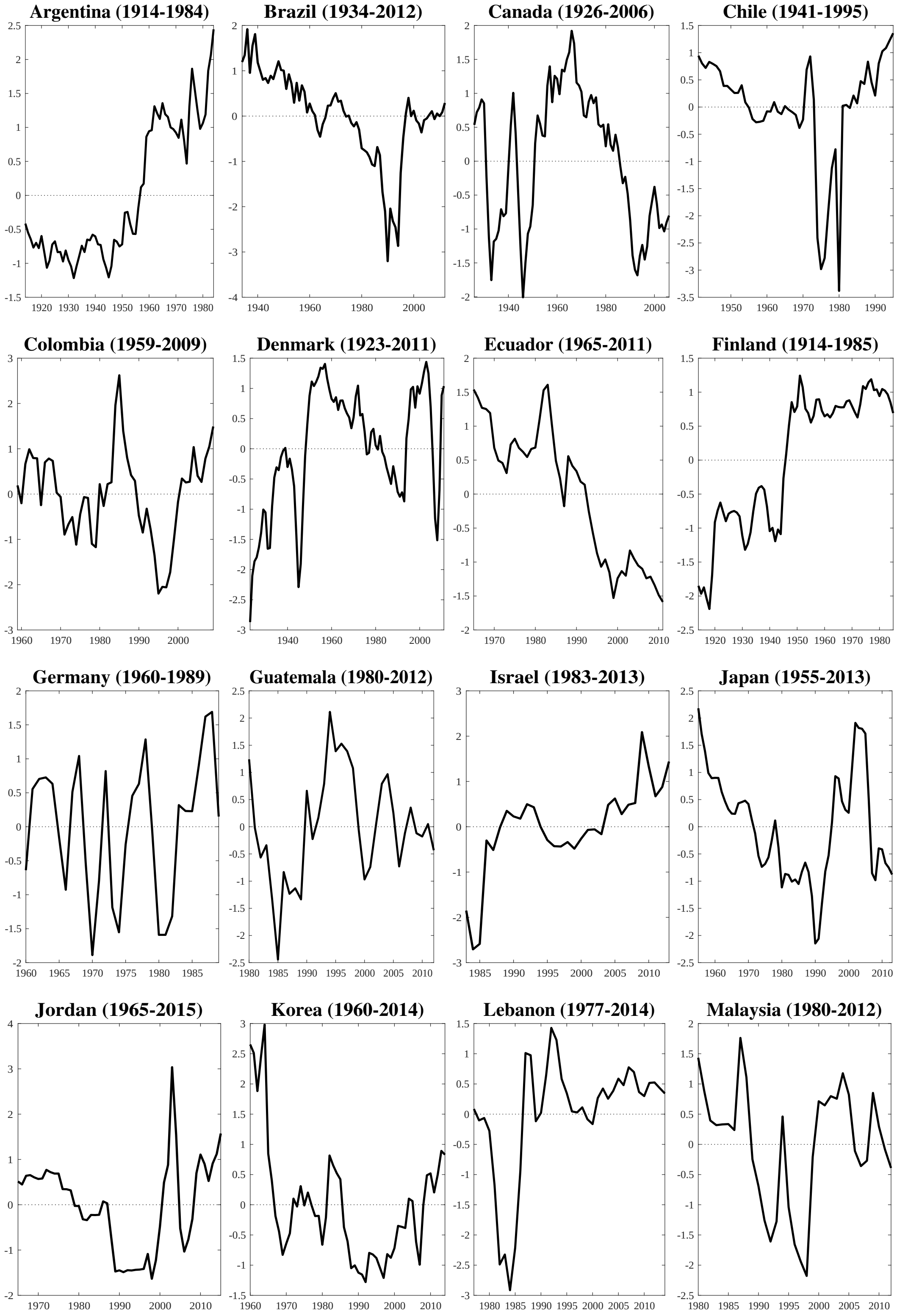

Figure B.3: Cointegration residuals based on the log-log specification

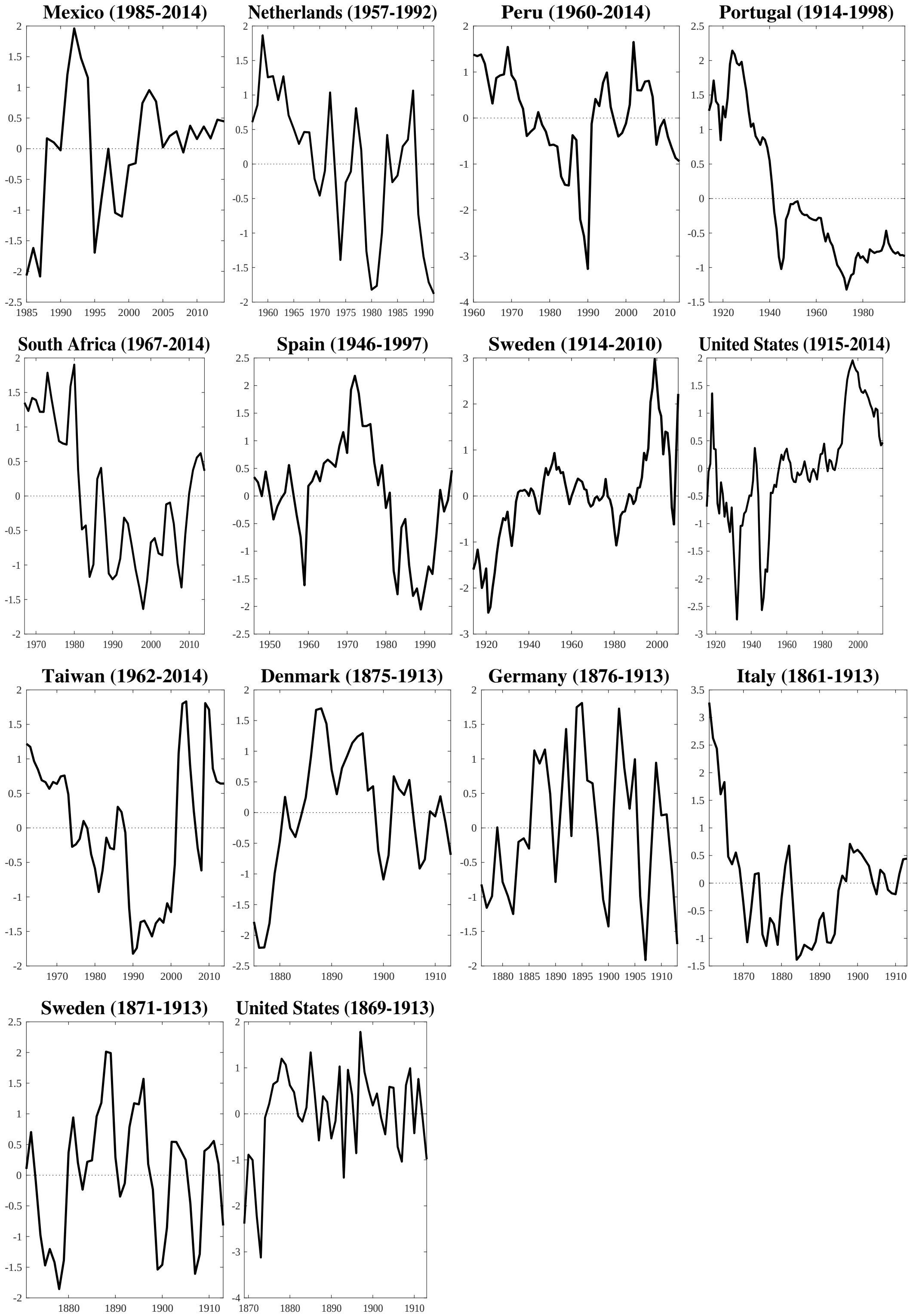

Figure B.4: Cointegration residuals based on the log-log specification

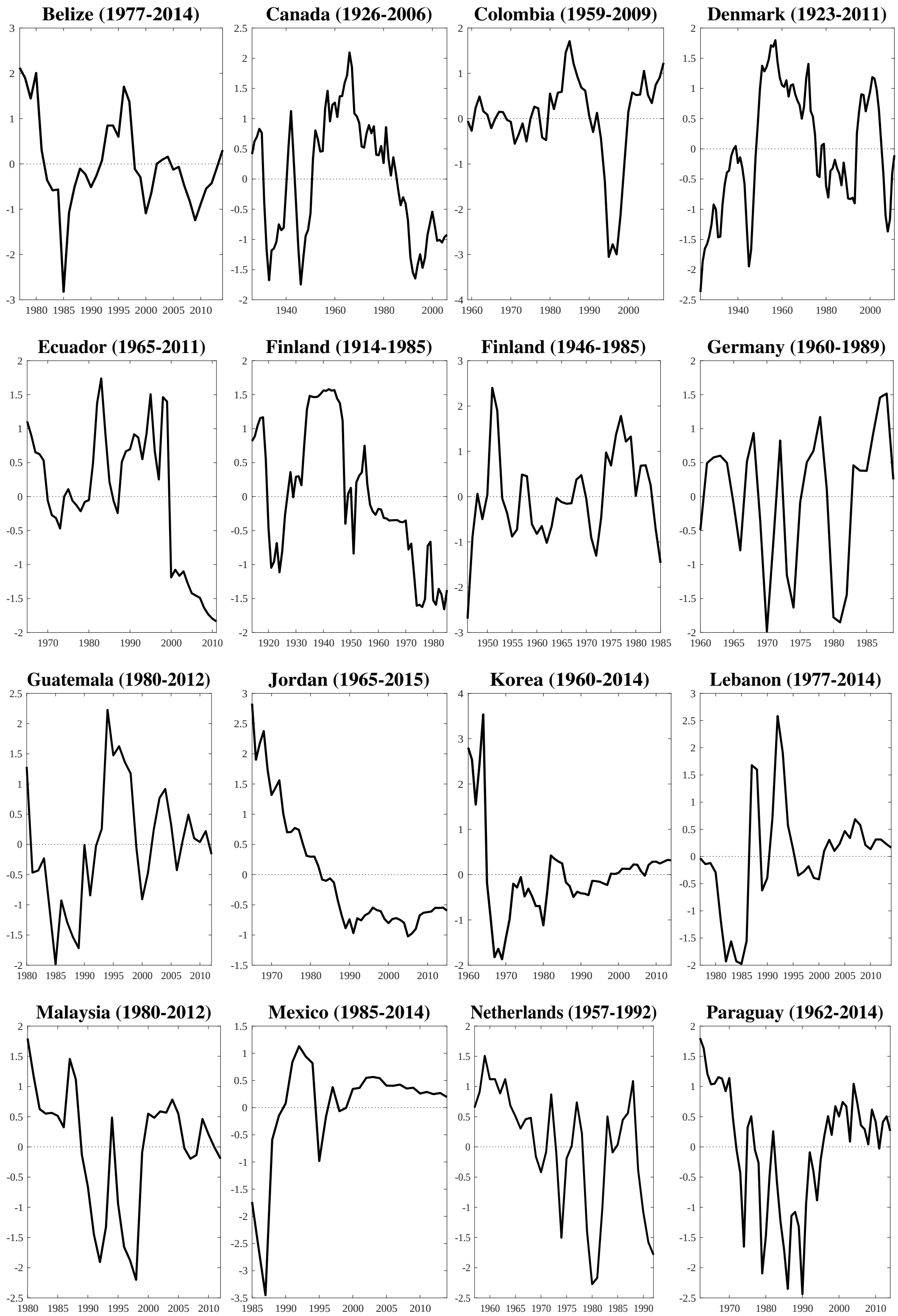

Figure B.5: Cointegration residuals based on the Selden-Latané specification

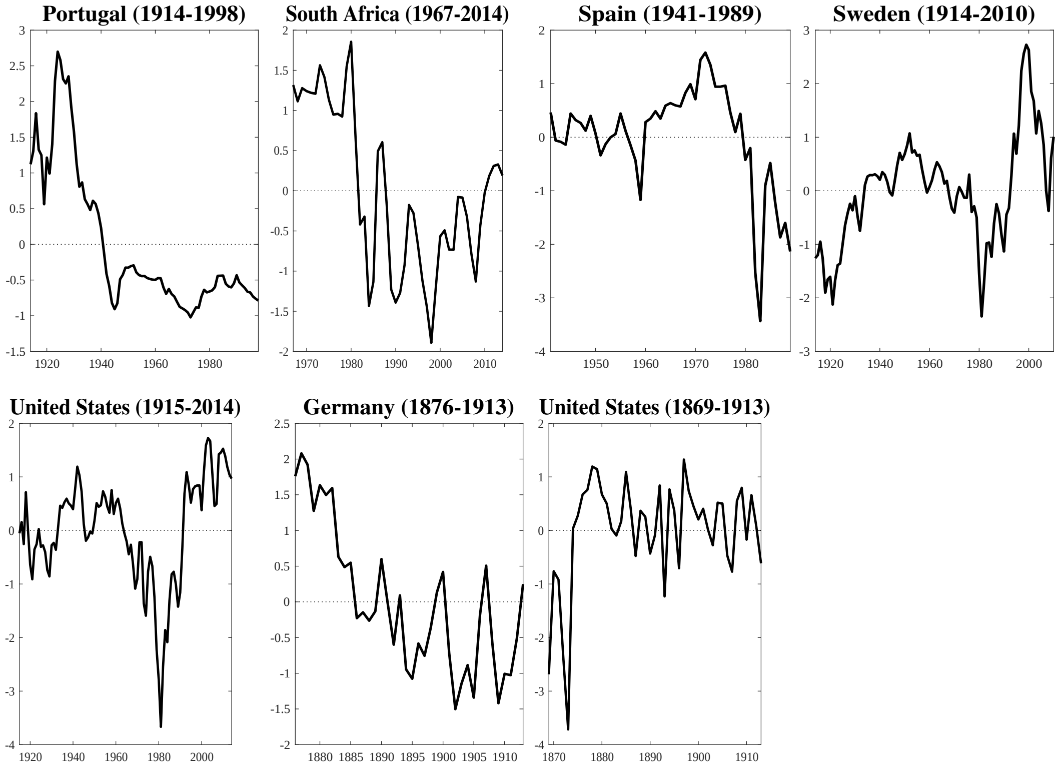

Figure B.6: Cointegration residuals based on the Selden-Latané specification

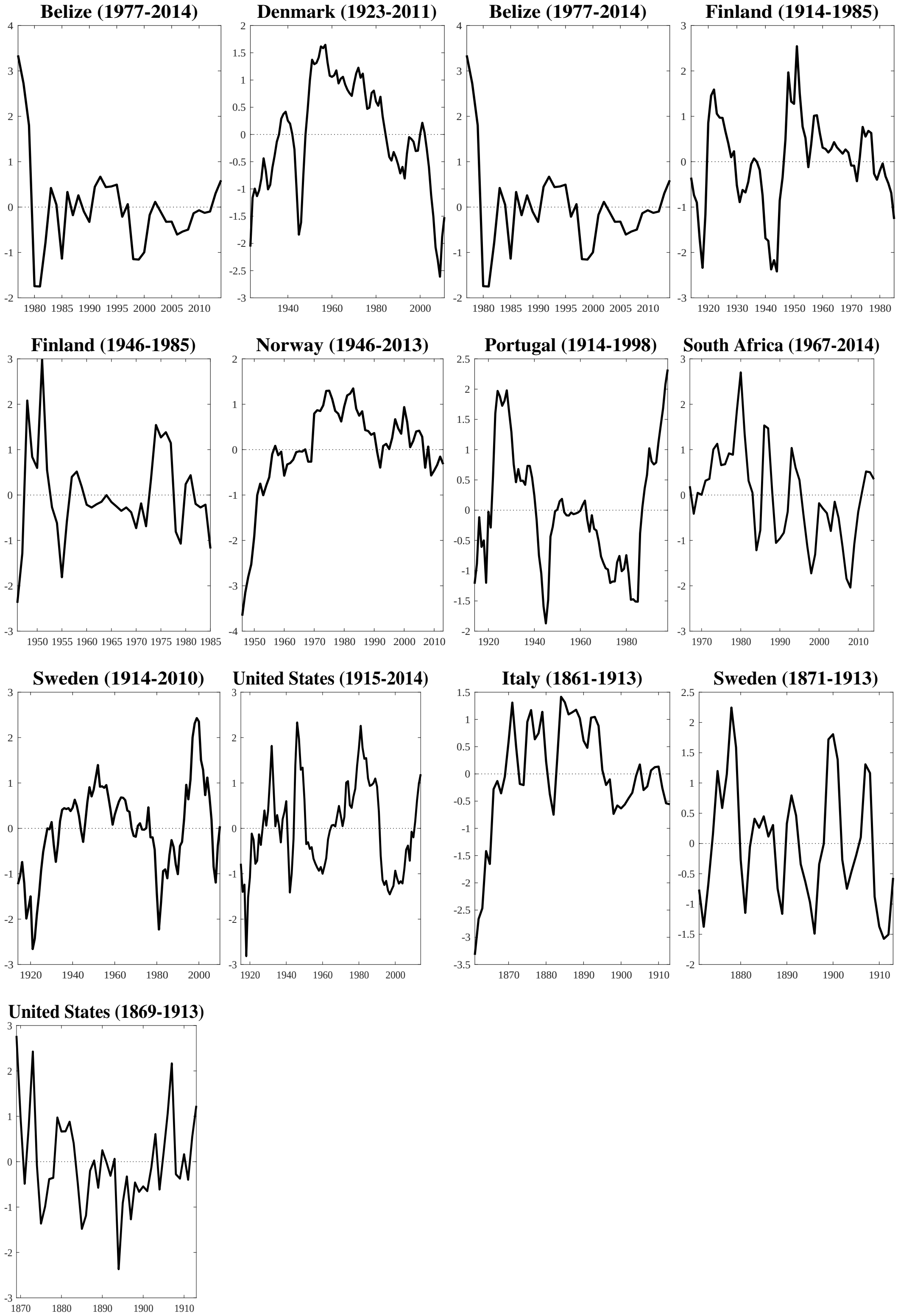

Figure B.7: Cointegration residuals based on the trivariate specification featuring nominal M_2, nominal GDP, and the short rate.

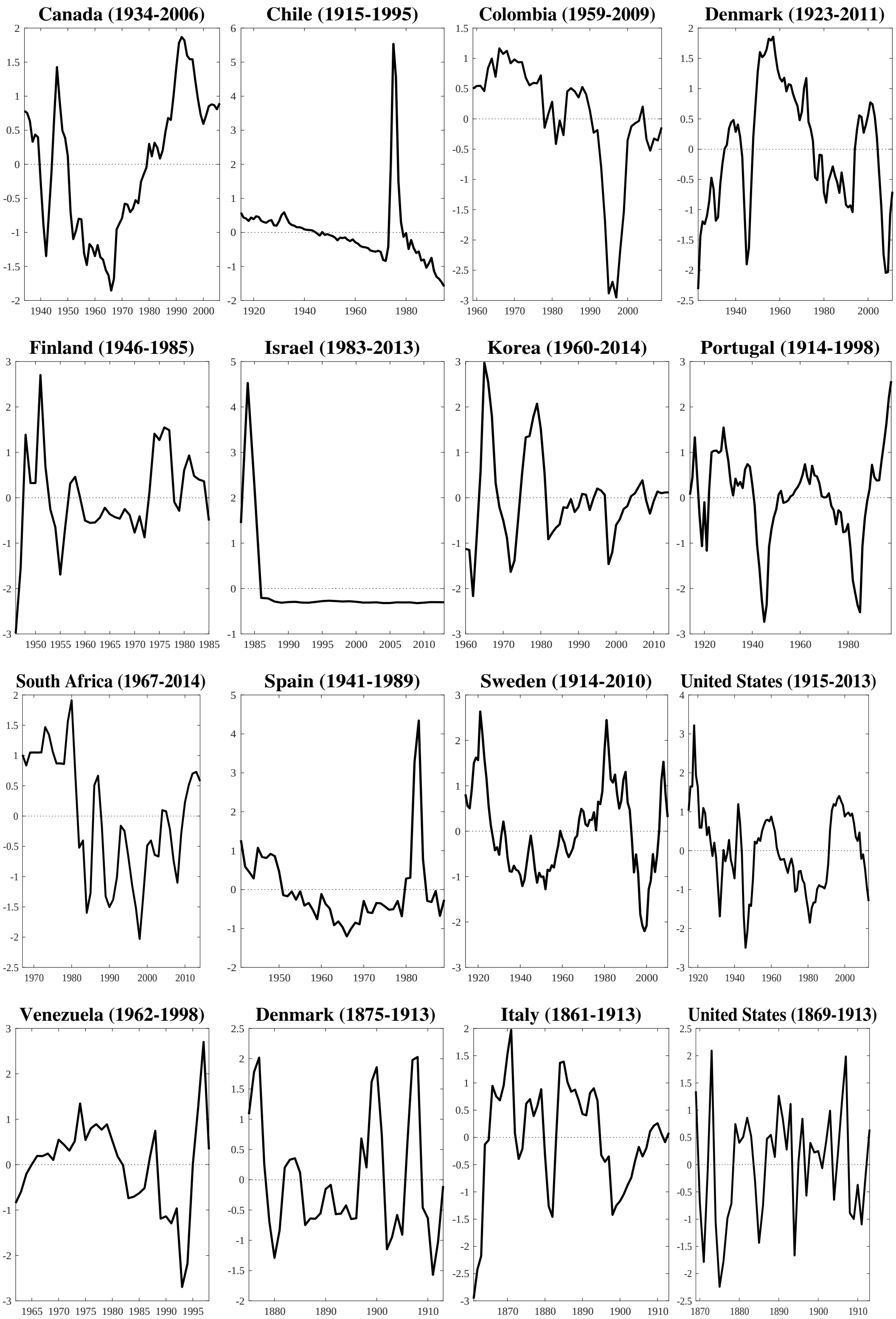

Figure B.8: Cointegration residuals based on the trivariate specification featuring real M_2, real GDP, and the short rate.

C United States with the Lucas-Nicolini M_2 aggregate

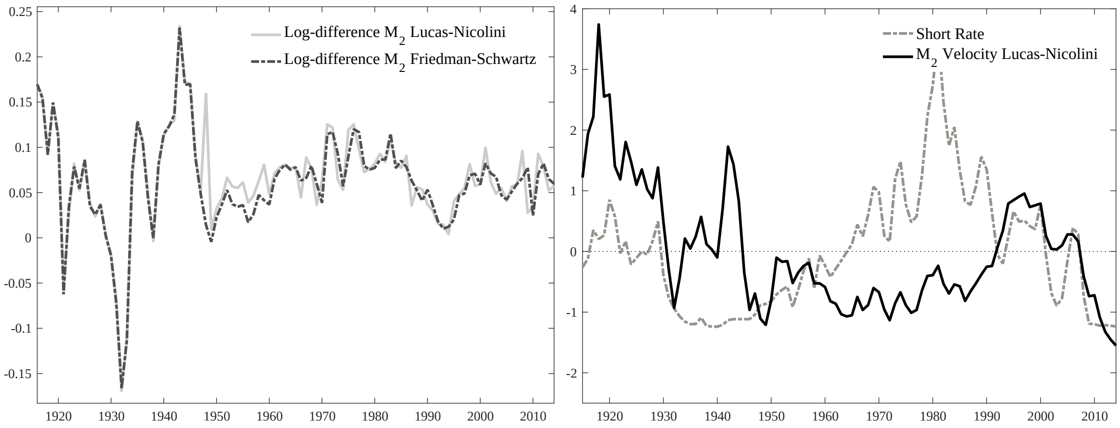

Figure C.9: Left-hand side: Log-differences of the Lucas-Nicolini M_2 series and the Friedman-Schwartz M_2 series; Right-hand side: M_2 velocity (Lucas-Nicolini M_2 aggregate) and the short rate

Based on the Lucas-Nicolini M_2 aggregate

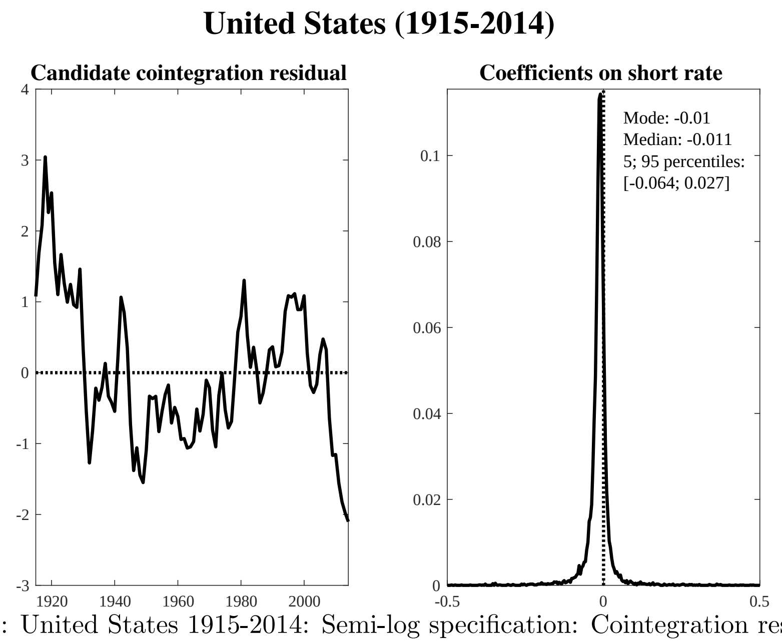

Figure C.10: United States 1915-2014: Semi-log specification: Cointegration residuals (left) and bootstrapped, bias corrected distributions of the coefficient of the short rate (right)

D Data

The data I use is largely based on the dataset created by Benati et al. (2018), which was kindly provided by Luca Benati. Almost all of the data are from original sources, i.e. either original hard copy, or central banks' or national statistical agencies' websites. In some cases the data of interest could not be found in original documents. In those cases the data are either from the *International Monetary Fund*'s (*IMF*) *International Financial Statistics* (*IFS*), the *World Bank*, or a dataset by Rolnick and Weber (1997) (the dataset is available at the 'WarrenWeber Collection' at the website of the Federal Reserve Bank of Minneapolis). The sources of the data series not coming from their dataset are reported in the table below.

Table D.5: Data sources for the additional series in the dataset

Country	Series and time period	Data source	Details
Argentina	M_2, 1884-1994	RW dataset	
Belize	M_2, 1977-2014	CB of Belize	
Brazil	M_2, 1901-1987, 1960-2012	RW dataset, World Bank	linked via splicing in 1960
Colombia	M_2, 1955-2012	IFS	
Denmark	M_2, 1975-1921, 1923-2011	Abildgren (2006), Abildgren (2012)	for the period 1975-1921: 'broad money'
Finland	M_2, 1866-1985	Haavisto (1992)	
Germany	M_2, 1876-1923, 1948-1992	RW dataset	
Guatemala	M_2, 1980-2014	Banco de Guatemala	'M2 Medios de Pago-Millones de quetzales'
Israel	M_2, 1982-2014	Bank of Israel	
Italy	M_2, 1948-1998	Banca d'Italia	
Japan	M_2, 1955-2013	IFS	
Jordan	M_2, 1964-2015	CB of Jordan	
Korea	M_2, 1960-2014	CB of Korea	
Lebanon	M_2, 1977-2014	Banque du Liban	
Malaysia	M_2, 1969-2012	Bank Negara Malaysia	Table 1.3 of the Monetary Aggregates
Portugal	M_2, 1854-1998	Banco de Portugal	Table 5
South Africa	M_2, 1967-2014	South African Reserve Bank	series code: KBP1373J
Spain	M_2, 1874-1997	López et al. (2005)	Cuadro 9.16 'Agregados Monetarios, 1865-1998'
Sweden	M_2, 1871-1959	Jonung (1975)	
	M_2, 1948-1959, 1960-2014	IFS, World Bank	linked via splicing
	GDP, 1560-2010	Schön and Krantz (2015)	nominal and real GDP
	short rate, 1856-2002	Historical Statistics Sweden[a]	Discount rate
	short rate, 2002-2014	Sveriges Riksbank	Reference rate
Taiwan	M_2, 1962-2014	CB of Taiwan	
Turkey	M_2, 1977-2014	CB of Turkey	
United States	M_2, 1869-1908	Balke and Gordon (1986)	
	M_2, 1909-1958, 1959-2014	Friedman and Schwartz (1970), FRED	
	M_2, 1915-2014	Lucas and Nicolini (2015)	1915-1947: Friedman and Schwartz (1970); 1947-1958: Rasche (1990); 1959-2014: FRED
Venezuela	M_2, 1940-1999	Banco Central de Venezuela	

[a] `http://www.historicalstatistics.org/htmldata15/index.html`