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Does Money Growth Predict Inflation?

Evidence from Vector Autoregressions Using Four Centuries of Data

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Does Money Growth Predict Inflation? Evidence from Vector Autoregressions Using Four Centuries of Data*

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Abstract

In this paper, we add new evidence to a long-debated macroeconomic question, namely whether money growth has predictive power for inflation or, put differently, whether money growth Granger causes inflation. We use a historical dataset – consisting of annual Swedish data on money growth and inflation ranging from 1620 to 2021 – and employ state-of-the-art Bayesian estimation methods. Specifically, we employ VAR models with drifting parameters and stochastic volatility which are used to conduct analysis both within- and out-of-sample. Our results indicate that the within-sample analysis – based on marginal likelihoods – provides strong evidence in favour of money growth Granger causing inflation. This strong evidence is, however, not reflected in our out-of-sample analysis, as it does not translate into a corresponding improvement in forecast accuracy.

JEL Classification: E31, E37, E47, E51, N13

Keywords: Time-varying parameters, Stochastic volatility, Out-of-sample forecasts

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1. Introduction

Whether monetary aggregates have predictive power for inflation is a key question in macroeconomics. But while the importance of monetary aggregates for inflation at a general level seems uncontroversial, views are more mixed when it comes to the narrower empirical relevance of monetary aggregates for forecasting inflation and/or the conduct of monetary policy.[1] These mixed views are also reflected in the world of central banking where the European Central Bank (ECB) – unlike most other central banks – ever since it was created has had a strong focus on analysis related to money in its policy framework; see, for example, ECB (2003) and Gerdesmeier (2009). The limited interest for money among other central banks is clearly related to the fact that models without money – such as mainstream DSGE models – have become the dominating tools both in research and at policy institutions. For instance, Woodford (2008, p. 1561) concluded that in the New Keynesian model, there is no "*compelling reason to assign a prominent role to monetary aggregates in the conduct of monetary policy*."

However, in the last decade, there has been an increased interest in money in the macroeconomic debate. The main cause of this revival is likely the wide range of non-standard monetary policy measures that central banks around the world used in order to support the banking system with liquidity and stimulate the macroeconomy around and after the financial crisis of 2008 and the following debt crisis in the euro area. The recent surge in inflation – which has substantially affected many countries around the globe – has further intensified this interest.[2]

The goal of this paper is to provide further empirical evidence concerning the predictive power that money growth has for inflation. We do this by analysing whether money growth Granger causes inflation in vector autoregressive (VAR) models. Granger causality is a concept that directly relates to predictive ability. In the context of this paper, money growth Granger causes inflation if a VAR where inflation is determined by lagged values of both inflation itself and money growth generates better forecasts than a VAR in which inflation only is determined by lagged values of itself. While Granger causality is a statement about predictive ability, it can be investigated based on both out-of-sample analysis and within-sample analysis. There are arguments in the literature in favour of both approaches and to provide as much evidence as possible on this topic, we will do both.

[1] See, for example, Lucas (1980), Thoma (1994), Nelson (2003), Bachmeier *et al.* (2007), Hale and Jordà (2007), Assenmacher-Wesche and Gerlach (2008), Sargent and Surico (2011), Gertler and Hofmann (2018) and Borio *et al.* (2023).

[2] It can be noted that the former governor of the Bank of England, Mervyn King has stated that "*Money has disappeared from modern models of inflation. You do not have to believe that there is a stable mechanical link between a particular measure of money and inflation to regret that development.*" (King, 2021, p. 4)

More specifically, we conduct our analysis using Swedish data on money growth and inflation which stretches over four centuries, from 1620 to 2021. We estimate bivariate VAR models for money growth and inflation in order to assess Granger causality. We employ models which allow for time-varying parameters and stochastic volatility – two features which seem necessary given the nature of the data. Relying on a very long sample, it is highly likely that the data-generating process has not been stable during the entire time period; one reason to suspect instability is of course the fact that there have been several different monetary regimes. Moreover, simply by looking at the data, one can argue that allowing for some type of heteroscedasticity seems essential (see Figure 1 in Section 3).[3] The models are estimated using state-of-the-art Bayesian methods developed by Chan and Eisenstat (2018a, 2018b). Relying on these methods, we not only have appropriate and advanced forecasting tools for the out-of-sample analysis, we also have an appealing way to assess the question of Granger causality based on within-sample analysis through formal model selection based on marginal likelihoods.

In conducting this analysis, we combine the use of historical macroeconomic data with modern analysis tools. We accordingly contribute to the literature on economic history as well as that on macroeconomics. The study can be seen as fairly unique. One aspect of this is our use of money-supply data; although series of inflation exist for several countries back to the Middle Ages, data on money supply covering four centuries are largely missing for most countries and analysis similar to ours is accordingly difficult to conduct. It can, however, be noted that formal econometric analysis of very long time series appears to be a topic of growing interest. Recent studies include Golez and Koudjis (2018) who – employing four centuries of data – analysed the predictability of stock returns, and Plakandaras *et al.* (2022) who studied real interest rates using time-varying parameter VAR models and data covering 700 years.

Briefly mentioning key findings, we note that the within-sample analysis provides strong evidence in favour of money growth Granger causing inflation. This strong evidence is, however, not reflected in the out-of-sample analysis as it does not translate into a corresponding improvement in forecast accuracy.

The remainder of this paper is organised as follows: In Section 2, we provide a brief literature review. We describe the data that we rely upon in Section 3. In Section 4, we present the methodological framework that we employ. In Section 5, we describe the empirical analysis and present our results. Finally, Section 6 concludes.

[3] There is also a growing literature in macroeconomics in general pointing to the importance of allowing for time variation in the dynamics of the model and/or the volatility of its disturbances; see, for example, Cogley and Sargent (2005), Primiceri (2005), Stock and Watson (2012), Abbate *et al.* (2015), Prieto *et al.* (2016), Akram and Mumtaz (2019) and Karlsson and Österholm (2020a, 2020b, 2023).

2. Literature review

In the history of economic thought there has been a long debate on what determines price fluctuations. Already in 1517, Copernicus formulated a type of quantity theory of money (Volckart, 1997, p. 430). The monetarist view is that money supply, divided by output, is the fundamental determinant of prices in the long-run.[4] Illustrating the monetarist view, we have Friedman's (1963, p. 17) famous statement that "*inflation is always and everywhere a monetary phenomenon*". The non-monetarist often point out that real factors, such as population growth or structural transformation, provide important explanations of many inflation episodes.

The mechanism behind inflation is, of course, different for commodity money – which historically have mainly consisted of silver and gold coins – and fiat money. A very large increase in the supply of money may have a similar effect under both commodity and fiat money. However, the amount of fiat money can be increased much more than commodity money, since the supply of the latter is ultimately restricted by the cost of production of precious metals, provided the mint equivalent is remained unchanged. Nevertheless, under commodity money, substantial increases in money supply are still possible if the mint equivalent is substantially increased, but then coins are more resembling fiat money – that is, they circulate periodically above their value as precious metal. Debasement as a source of inflation is a growing research field (Sussman, 1993; Velde *et al.*, 1999; Redish, 2000; Sussman and Zeira, 2003; Edvinsson, 2011), but is limited by the lack of long time series on both inflation and money supply.

Schwartz (1973) analysed the major inflation periods since antiquity, finding support for the monetarist view, for both commodity and paper money. When money supply increased, there was a corresponding increase in prices, and this applied to both commodity and fiat monies; the only exception to this was the Great Debasement in 16th century England.

Among economic historians, there has been a long debate whether the silver price inflation – that is, the decrease in the purchasing power of silver coins of stable fine metal content – following Europe's discovery of America was mainly caused by the inflow of silver or population growth. The monetarist view is that it is almost self-evident the cause was the inflow of silver. However, historians have argued that there is a discrepancy in timing; while inflation mainly took place in the 1530s and 1540s, the inflow of silver only accelerated after the 1550s. There was also a corresponding increase in the purchasing power of silver following the Black Death, and the purchasing power of silver at the end of the 16th century only seems to have reverted to previous levels. An alternative view is that price inflation followed population growth, which put upward

[4] This rests on an assumption of a relatively stable velocity.

pressure on agricultural prices (Golstone, 1984; Fischer, 1996). When the population increases, people sacrifice some liquidity in order to maintain consumption. Golstone (1984) also argues that there was a positive effect of urbanisation on prices, given that urbanisation increases demand for money. Recently Melitz and Edo (2019) argued that silver inflow and population growth both mattered separately and jointly.

Recently, the monetarist view has been questioned, given that money supply increased substantially after the Great Recession, but that inflation did not seem to be affected much by this. However, the recent surge in inflation during the recovery after the Covid pandemic appears to have opened up the possibility for a lively debate on this topic. An early overview and some empirical analysis can be found in Borio *et al.* (2023) who – while reluctant to draw strong conclusions – note that (p. 2) "*An upsurge in money growth preceded the inflation flare-up, and countries with stronger money growth saw markedly higher inflation*."

Regarding related literature, we finally note that our paper is similar to more contemporary research in macroeconomics and monetary policy that has focused on whether money growth has predictive power for inflation. Results from out-of-sample analysis is somewhat mixed. Several studies suggest that monetary aggregates can help improve inflation forecasts though. This includes Bachmeier and Swanson (2005) and Berger and Österholm (2011b) who used US data, and Hofmann (2009), Berger and Österholm (2011a) and Dreger and Wolters (2014) who relied on data for the euro area. However, it can also be noted that Berger and Österholm (2011a, 2011b) point out that the predictive power of money growth for inflation tends to be substantially lower in both the United States and the euro area in more recent sample periods – approximately since the late 1980s – compared to the 1970s and early 1980s. This indicates that the predictive power of money growth for inflation might vary over time – for example due to different regimes – something which also Borio *et al.* (2023) suggest. The issue of different regimes is something that we turn to next in the present paper as we present our historical data.

3. Data

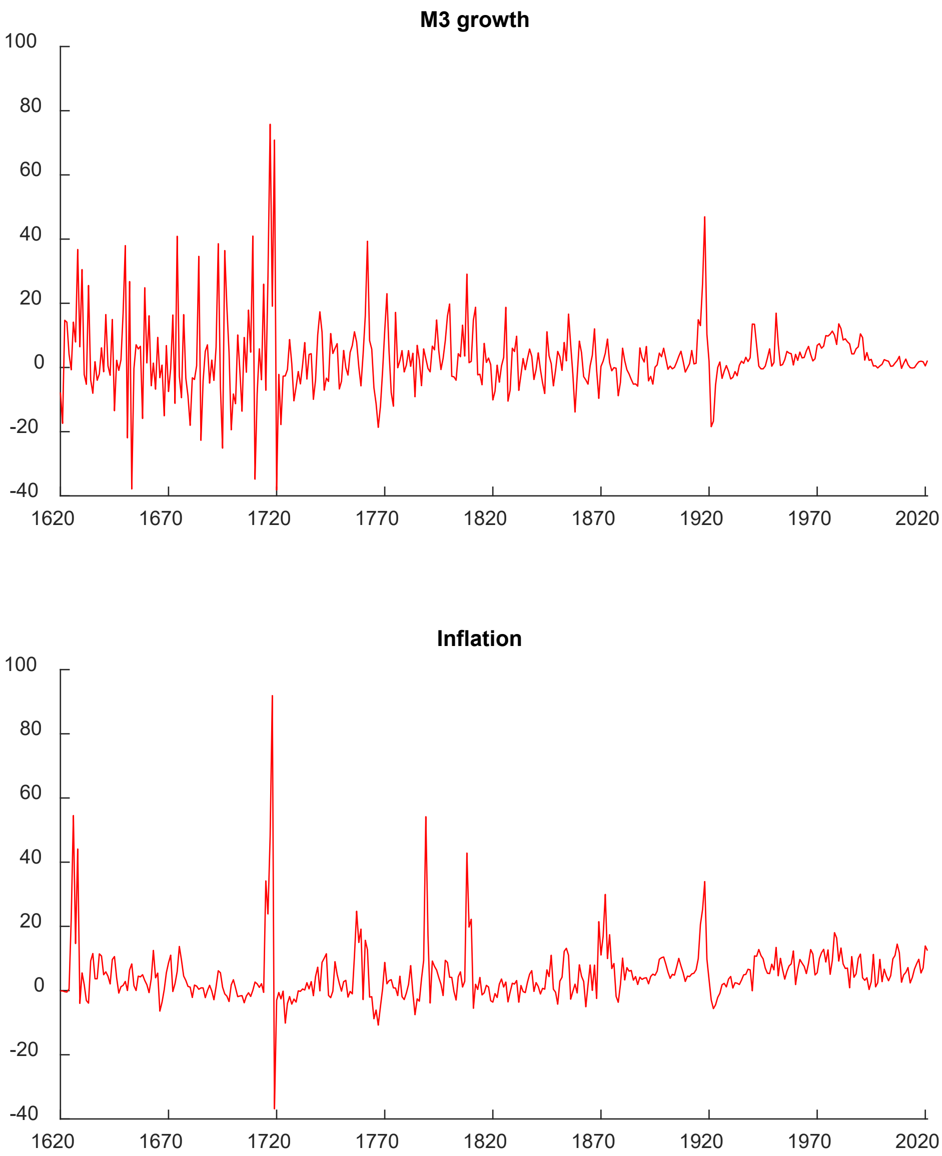
We use two data series in our analysis: prices and money supply. From 1914 to 2021, prices are given by the official series – as published by Statistics Sweden – of the consumer price index. Before 1914, the price series is the consumer price index from Edvinsson and Söderberg (2010, 2011); this has been calculated using prices of a basket of goods and services, whose composition shifted somewhat over time. The period consists of several monetary regimes, alternating between metallic and paper standards. The metallic standards were based on silver up to 1624, copper and silver 1624-1776 (that is, a type of bimetallism), silver 1776-1873, and gold from 1873 until the fall of the Bretton Woods system. However, even before 1971, under several periods the convertibility of the dominant currency into coins of intrinsic value was suspended. Up to the early 19th century, at periods, several different currencies were used in Sweden, with a floating exchange rate relative to each other. For example, in the 1740s, five different metallic currencies were used alongside inconvertible paper notes. The consumer price index follows the weakest currency, which also tended to be the dominant means of payment.

The historical data on money supply reflect M3 and are taken from Edvinsson and Ögren (2014), which are linked to financial data from Statistics Sweden for the latest years (from 2013 to 2021 to be specific). For the period before 1819, data on bank deposits are lacking, but they were quite small; therefore, the M3 series is constructed by adding private bank notes to M0 for the period 1804-1818, while before 1804, M3 is assumed to be the same as M0. Before 1804, M0 is composed of means of payments expressed in different currency units, which are transformed into the same unit as the consumer price index, using the market exchange rates between those currency units. The first paper notes circulated in 1661, while before that year the money supply entirely consisted of intrinsic value coins. Paper money dominated money supply from the mid-18th century. In our empirical analysis, we model the annual percentage change in the variables. Data on CPI inflation and money growth are shown in Figure 1.

Up to 1624, Sweden was – as mentioned above – on a silver standard. The silver coins were commodity, entailing that the value of the coins could be guaranteed by their silver content. Excessive minting of silver coins without debasement could not cause long-term inflation, since the decline in the value of silver coins below the bullion point would cause the melting down of such coins, or the export of them as bullion. The only mechanism by which increased money supply could have caused inflation was the debasement of coins; rapid debasement occurred on several occasions during the 16th century (Edvinsson, 2011). However, the older silver coins would still retain their value under such conditions, which can be contrasted to modern inflation, whereby all means of payments functioning as money and denominated in a unit of account fall in value.

Sweden has had some form of fiat, or at least semi-fiat, money since 1624 – that is, the value of money was not guaranteed by the intrinsic properties of the means of payment. In 1624, a copper standard was introduced in Sweden, which existed in parallel with the silver standard; this system existed until 1777 when the silver standard was reintroduced. The copper coins were a hybrid between commodity and fiat money. Sweden was a copper monopolist. The idea behind introducing copper coins was to mint such when the international copper price stood low. In this way, copper was withdrawn from the international copper market which put an upward pressure on the copper price (Edvinsson, 2012). This system was based on setting the value of copper coins substantially above their intrinsic metal value. Therefore, usually there was no incentive to export or smuggle copper coins, given that their value as coins was much higher than their value as bullion for export. This also implied that a large volume of copper minting did not cause to the melting down or export of copper coins, but rather to the fall in their value as money – that is, inflation. For example, large minting in the 1620s led to the devaluation of copper coins by 50 percent in 1633, relative to silver coins. This created a complicated monetary system with five different metallic currencies, at period with a floating exchange rate relative each other. Due to the high transaction costs when using copper money, the first paper notes in Europe were introduced in 1661 by Stockholm Banco, which caused a panic and the reorganisation of the bank into Sveriges Riksbank, the world's oldest central bank. In the 1710s, copper tokens functioning as fiat money led to substantial inflation, while later during the 18th century paper money came to dominate money supply, which also caused inflation.

Figure 1. Data



Note: Both variables are measured in percent.

When the silver standard was reintroduced in 1777, with the riksdaler as the main unit of account, most transactions were probably made in paper money. In 1789 the king decided to supersede Sveriges Riksbank, when the bank refused to lend him money to wage war with Russia, and the king issued his own paper notes, later denominated as riksdaler riksgälds. Up to around 1820, the riksdaler riksgälds fell in value by 75 percent relative the riksdaler specie coin. Afterwards, monetary stability followed, up to 1914. In 1873 Sweden switched to a gold standard. During the course of the 19th century bank money expanded sharply its share of M3. Private banks also issued their own paper notes convertible into Riksbank money, although they were not legal tender. These notes continued to circulate until the Riksbank gained monopoly on the issuing of banknotes in 1904. Before the industrial breakthrough in the late 19th century, short-term fluctuations in consumer prices were largely induced by harvest fluctuations, which were exogenous supply shocks, while this effect disappeared during the course of the 20th century. Prices could therefore fluctuate sharply even when the money supply remained constant.

Although Sweden was on a gold standard for the whole period from 1873 to the fall of the Bretton Woods system in 1971, there were two periods when convertibility into gold was suspended – in 1914-1922 due to the First World War and 1931-1951 following the Great Depression. From a historical point of view, the system after 1971 is very different. Up to 1971, the value of money was always expressed in precious metal – whether it be gold, silver or copper – but from then on money lost all its connection to precious metal, which also impacted on the relation between money supply and inflation. The inflation targeting that was introduced in Sweden in 1995 could not have been pursued in the pre-industrial economy, during which stability of the value of money was defined relative to precious metal.

This historical overview suggests that the dynamics between money supply and inflation may have been different over time. In our empirical analysis, we accordingly employ models that allow for this and also have a closer look at the out-of-sample forecast results for four sub-periods. The periodisation corresponds to the succession of different monetary standards in Sweden: the copper and silver standard 1624-1776, the silver standard 1777-1873, the gold standard 1873-1971, and the post-Bretton Woods system in the last decades. However, also during these sub-periods, extreme events occurred with different dynamics; these include the dramatic fall of the purchasing power of copper coins 1626-1628, the massive minting of inconvertible coin tokens 1715-1719, the spread of inconvertible riksgälds notes by the king competing with the notes of the Riksbank 1789-1804, the collapse of the value of Riksbank notes in 1808-1810 due to suspension of convertibility into silver coins, and the suspension of the gold standard during World War I.

4. Methodological framework

In this section, we present the tools we use for our empirical analysis. We first describe the Bayesian VAR models and then how we conduct model selection in order to draw conclusions regarding Granger causality within-sample. We then explain our out-of-sample forecast exercise and how Granger causality is assessed based on it.

4.1 The Bayesian VAR models

The models used for the analysis conducted in this paper will be bivariate VAR models which are estimated using Bayesian methods. Given the nature of the data, it seems essential to allow for heteroskedastic error terms in the model; we here choose to model this as stochastic volatility. In addition, it seems likely that the dynamic relationship between inflation and money growth has changed over time; we allow for this possibility by using a specification with time-varying parameters, where the time variation is modelled as the parameters drifting over time. We accordingly estimate two different specifications of the VAR:

i) Constant parameters and stochastic volatility (henceforth denoted "SV")

ii) Time-varying parameters and stochastic volatility (henceforth denoted "TVP-SV")

Based on the marginal likelihoods of the two models – an issue which we will return to below – we can then assess the empirical relevance of the time-varying parameters.

Presenting the framework – which is based on Chan and Eisenstat (2018a) – we initially specify the most general version of the VAR, namely the TVP-SV model; the SV model is a restricted version of the TVP-SV model. The vector of dependent variables is given by $\boldsymbol{y}_t = (\pi_t \quad \mu_t)'$, where π_t is the CPI inflation rate (in percent) and μ_t is the growth rate of money (in percent). Equation (1) presents the model:

$$\boldsymbol{B}_{0t}\boldsymbol{y}_t = \boldsymbol{\delta}_t + \boldsymbol{B}_{1t}\boldsymbol{y}_{t-1} + \cdots + \boldsymbol{B}_{pt}\boldsymbol{y}_{t-p} + \boldsymbol{\varepsilon}_t \qquad (1)$$

Turning to the components of the equation, $\boldsymbol{B}_{0t}$ is a 2x2 lower triangular matrix with ones on the diagonal. $\boldsymbol{\delta}_t$ is a 2x1 vector of time-varying intercepts and $\boldsymbol{B}_{1t}, \dots, \boldsymbol{B}_{pt}$ are 2x2 matrices with the parameters describing the dynamics of the model. We set the lag length in our estimations equal to $p = 2$, which is a standard choice in the literature; see, for example, Cogley and Sargent (2005), Primiceri (2005) and Aastveit *et al.* (2017). $\boldsymbol{\varepsilon}_t$ is a 2x1 vector of disturbances, $\boldsymbol{\varepsilon}_t \sim N(\boldsymbol{0}, \boldsymbol{\Sigma}_t)$, where $\boldsymbol{\Sigma}_t = diag\left(exp(h_{\pi t}), exp(h_{\mu t})\right)$. The model in equation (1) is hence written on a "structural" form – that is, the disturbances $\boldsymbol{\varepsilon}_t$ are uncorrelated.

The fact that the model is written on a structural form is related to $\boldsymbol{B}_{0t}$ being lower triangular. It can be noted that this assumption implies that inflation can affect money growth contemporaneously – that is, within the year – but that money growth only can affect inflation with a lag. We argue that this is the most intuitive specification – at least from a present-day monetary-policy perspective (if money is seen as the policy instrument). Needless to say, using annual data, this restriction is unlikely to hold; while prices might be sticky, they are unlikely to not respond at all within a year. There may also be differences between the 20th century and earlier periods. Supply shocks in earlier periods due to harvest failures increased prices in the same year without that leading to an expansion of the money supply, while monetary shocks usually affected prices quite quickly. For example, the growth of coin tokens earlier in 1718 most likely contributed to price increases towards the end of the year. Fortunately, this restriction is not crucial for our Granger-causality analysis. In our in-sample analysis, the coefficients in $\boldsymbol{B}_{1t}, \dots, \boldsymbol{B}_{pt}$ are the key components (which we explain in Section 4.3 below). And in our out-of-sample analysis, we rely on the model's reduced form when generating forecasts.[5] So even though the structural form of the model may have a property that can be questioned, this is not a fundamental flaw. It should also be noted that we want to work with the model in its structural form since this is an important part of simplifying the numerical evaluation of the model and accordingly is a central feature in allowing us to make our formal model comparisons.

Turning to the log-volatilities, we model these as random walks:

$$\boldsymbol{h}_t = \boldsymbol{h}_{t-1} + \boldsymbol{\zeta}_t \tag{2}$$

where $\boldsymbol{\zeta}_t \sim N(\boldsymbol{0}, \boldsymbol{\Sigma}_{\boldsymbol{h}})$. Finally, the free parameters of $\boldsymbol{\delta}_t$ and $\boldsymbol{B}_{it}$ are gathered in the parameter vector $\boldsymbol{\theta}_t$, whose evolution also is specified as a random walk:

$$\boldsymbol{\theta}_t = \boldsymbol{\theta}_{t-1} + \boldsymbol{\eta}_t \tag{3}$$

where $\boldsymbol{\eta}_t \sim N(\boldsymbol{0}, \boldsymbol{\Sigma}_{\boldsymbol{\theta}})$. Note that when the SV model is estimated, we impose the restriction that $\boldsymbol{\Sigma}_{\boldsymbol{\theta}}$ is zero; this forces the parameters in $\boldsymbol{\theta}_t$ to be constant during the entire sample.

The choice of prior distribution and information is an important part of Bayesian inference. Here we aim to be relatively uninformative and, as we are comparing different versions of the model, not use prior infor-

[5] We get the reduced form of the model by pre-multiplying the model in equation (1) with $\boldsymbol{B}_{0t}^{-1}$. That is, the reduced form is given as $\boldsymbol{y}_t = \boldsymbol{\delta}_t + \boldsymbol{A}_{1t}\boldsymbol{y}_{t-1} + \cdots + \boldsymbol{A}_{pt}\boldsymbol{y}_{t-p} + \boldsymbol{e}_t$ where $\boldsymbol{\delta}_t = \boldsymbol{B}_{0t}^{-1}\boldsymbol{\mu}_t$, $\boldsymbol{A}_{it} = \boldsymbol{B}_{0t}^{-1}\boldsymbol{B}_{it}$ and $\boldsymbol{e}_t = \boldsymbol{B}_{0t}^{-1}\boldsymbol{\varepsilon}_t$.

mation that favours one particular specification. We need to specify priors for the initial state of the parameters, $\boldsymbol{\theta}_0$, and the log-volatilities, $\boldsymbol{h}_0$ and the variances of the change in the parameters, $\boldsymbol{\Sigma_\theta}$, and log-volatilities, $\boldsymbol{\Sigma_h}$. For $\boldsymbol{\theta}_0$ we use an uninformative normal prior $\boldsymbol{\theta}_0 \sim N(\mathbf{0}, 5\boldsymbol{I})$ and the prior for $\boldsymbol{h}_0$ is set to match the residual variance from an OLS-fit of a univariate AR(2) model. This is done by way of a normal prior, $h_{i,0} \sim N(\mu_i, 0.25)$ with μ_i chosen to set the expected value of $exp(h_{i,0})$ equal to the residual variance. For specifications with constant parameters this also serves as a prior for the parameters. The variance matrices $\boldsymbol{\Sigma_\theta}$ and $\boldsymbol{\Sigma_h}$ of the increments to the parameters and log-volatilities are assumed to be diagonal and we use inverse Gamma priors, $iG(v, S)$, for the diagonal elements with the shape parameter fixed at $v = 5$ and the scale parameter, or equivalently the prior mean of the variance, $S/(v-1)$, is chosen in an empirical Bayes fashion – that is, we use a grid search to find the parameter values that maximize the marginal likelihood or fit of the model.[6] We use different prior settings for the constant terms, remaining parameters in $\boldsymbol{\theta}$ and the volatilities with the grid search resulting in prior means of the variances of 0.04 for the constants, 0.0002 for other regression parameters in $\boldsymbol{\theta}$ and 0.36 for the log-volatilities.

As the posterior distribution and marginal likelihood of this model class are not available in closed form, we use Markov Chain Monte Carlo to simulate from the posterior distribution and estimate the marginal likelihood using the methods developed in Chan and Eisenstat (2018b)

4.2 Model selection

In order to establish which model is preferred by the data, we rely on Bayes factors for model selection. The Bayes factor is the posterior odds in favour of the null hypothesis when the prior probability of the null hypothesis is 0.5. Denoting the model under the null hypothesis as M_0 and the model we are comparing it to as M_A, we define the Bayes factor as

$$B_{0A} = \frac{m(\boldsymbol{y}|M_0)}{m(\boldsymbol{y}|M_A)} \tag{4}$$

where $m(\boldsymbol{y}|M_0)$ and $m(\boldsymbol{y}|M_A)$ are the marginal likelihoods of M_0 and M_A respectively, and $\boldsymbol{y}$ are the data. When assessing the strength of the evidence, we rely on the transformation $2ln(B_{0A})$ as is commonly done; see Kass and Raftery (1995) for an overview. In our empirical analysis, model selection will first be employed in order to establish which of the two models presented above is the most suitable with respect to modelling

[6] See Carlin and Louis (2002) for an introduction to empirical Bayes methods and Karlsson and Österholm (2020a) for analysis that also has adopted this approach.

the data. In a second step, we use it to establish whether money growth Granger causes inflation. We next turn to how the latter issue is done.

4.3 Within-sample analysis of Granger causality

A within-sample analysis of whether money growth Granger causes inflation is a fairly straightforward exercise. If money growth does not Granger cause inflation, this means that the model in equation (1) is characterized by a lower-triangular structure regarding its dynamics, that is, $\boldsymbol{B}_{1t}$ and $\boldsymbol{B}_{2t}$ are both lower triangular. (Recall that we have set lag length in the model to $p = 2$.) This implies that inflation is exogenous in the time series sense with respect to money growth.[7] Assessing the relevance of the restriction that the elements $\boldsymbol{B}_{1t}^{12}$ and $\boldsymbol{B}_{2t}^{12}$ (that is, the row 1, column 2 elements of $\boldsymbol{B}_{1t}$ and $\boldsymbol{B}_{2t}$) both are zero can be done by estimating two different models. The first is a fully endogenous model, in which inflation can be affected by lagged money growth; in this case, no restrictions are placed on $\boldsymbol{B}_{1t}$ or $\boldsymbol{B}_{2t}$. In the second model, the restriction that $\boldsymbol{B}_{1t}^{12} = \boldsymbol{B}_{2t}^{12} = 0$ is imposed so that inflation is exogenous with respect to money growth. The two models are then compared based on their marginal likelihoods.

4.4 Out-of-sample analysis of Granger causality

Relying on within-sample analysis to establish Granger causality has certain benefits. First, the concept comes with clear restrictions regarding the model structure – restrictions whose empirical relevance we can assess. Second, within-sample analysis tends to make efficient use of the all available data in a sample.[8] However, it is a well-known finding in the forecasting literature that conclusions based in-sample fit do not necessarily carry over to out-of-sample forecasting performance.[9] Seeing that Granger causality is a statement concerning predictability, there is accordingly a large literature that advocates out-of-sample analysis when investigating the issue of Granger causality empirically.[10]

In this paper we assess the issue of Granger causality out-of-sample by looking at the forecasting performance of the competing models. This is done through an out-of-sample forecast exercise where we estimate the competing models on an expanding sample. Each time we add an observation to the sample, we re-estimate the models and generate forecasts four years ahead. We return to the exact timing of this exercise below.

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[7] See Hamilton (1994, p. 303).

[8] See Diebold (2015).

[9] See, for example, Clark (2004) and Rossi and Sekhposyan (2011).

[10] An early important contribution is Ashley *et al.* (1980),

Forecasting performance is measured by the root mean square forecast error (RMSFE). This is defined as

$$RMSFE_h = \sqrt{\frac{1}{N_h}\sum_{i=0}^{N_h-1}(e_{t+h+i})^2} \qquad (5)$$

where $RMSFE_h$ is the RMSFE at horizon *h*. $e_{t+h+i} = y_{t+h+i} - \hat{y}_{t+h+i|t+i}$ is the forecast error, where y_{t+h+i} is the actual value of inflation at time *t+h+i* and $\hat{y}_{t+h+i|t+i}$ is the forecast made at time $t + i$. Finally, N_h is the number of forecast errors being evaluated at horizon *h*. If the model where inflation is endogenous with respect to money growth has a lower RMSFE at horizon *h* than the competing model (where inflation is exogenous with respect to money growth), we say that money growth Granger causes inflation at horizon *h*.

However, the RMSFEs are not population measures but should rather be seen as "point estimates". In order to take this fact into account – and test whether one model outperforms the other in a statistically significant manner – we also conduct a Diebold-Mariano test (Diebold and Mariano, 1995). Using this, we test the null hypothesis of equal forecasting ability between the two models. The test is conducted the following way: We denote the sequence of forecast errors at horizon *h* from the model in which inflation is endogenous with respect to money growth $\{e_{t+h+i}^{endo}\}_{i=0}^{N_h-1}$ and the corresponding forecast error for the model in which inflation is exogenous $\{e_{t+h+i}^{exo}\}_{i=0}^{N_h-1}$. We then calculate a sequence of differences according to $d_{h,t+h+i} = (e_{t+h+i}^{endo})^2 - (e_{t+h+i}^{exo})^2$ for each forecast horizon *h*. A test of equal forecasting ability of the two models is based on the test statistic

$$S_h = \frac{\bar{d}_h}{\sqrt{\frac{2\pi\widehat{f_d}(0)}{N_h}}} \qquad (6)$$

where $\bar{d}_h = N_h^{-1}\sum_{i=0}^{N_h-1} d_{h,t+h+i}$ and $\widehat{f_d}(0)$ is a consistent estimate of the spectral density at frequency zero. In practice, this is done by running a regression (for each forecasting horizon) where the dependent variable $d_{h,t+h+i}$ is explained by a constant alone. That is, we estimate

$$d_{h,t+h+i} = c + v_{h,t+h+i} \qquad (7)$$

where $v_{h,t+h+i}$ is an error term. If the constant c is found to be significant – based on a t-test using Newey-West standard errors and $N(0,1)$ critical values – we conclude that the null hypothesis of equal forecasting ability is rejected.[11]

5. Results

We first estimate the two different versions of the VAR specified in Section 4 – that is the TVP-SV and the SV models – in order to see which modelling choice best describes the data. The natural logarithm of the marginal likelihoods are presented in Table 1 and show that the TVP-SV model is preferred by the data. The evidence in favour of this model over the model with constant parameters is impressive. Using the TVP-SV model as the model under the null hypothesis, it is concluded – using the terminology of Kass and Raftery (1995) – that the evidence is "very strong" as $2ln(B_{0A})$ is as high as 103.6.[12]

Table 1. Log marginal likelihoods for VAR models with different assumptions regarding parameters.

	Log marginal likelihood
TVP-SV	-2602.3
SV	-2654.1

Note: Model estimated on data ranging from 1622-2021.

Having concluded that we should employ a model with drifting parameters and stochastic volatility – rather than a model with constant parameters and stochastic volatility – we next turn to the issue of Granger causality. This is first assessed using within-sample analysis.

5.1 Within-sample results regarding Granger causality

Results from this exercise are shown in Table 2. As can be seen, it is clear that money growth is judged to Granger cause inflation. If we denote the model in which inflation is endogenous M_0 and the model in which it is exogenous M_A – and again use the terminology of Kass and Raftery (1995) – the evidence is "very strong" as $2ln(B_{0A})$ is 17.6.

[11] Seeing that we are generating forecasts more than one step ahead from Bayesian VAR models which are nested and estimated on an expanding sample, the standard Diebold-Mariano test is not strictly valid here. However, Diebold (2015) points out that if the Diebold-Mariano test is used for model selection, one can use the setup that we employ here rather than more elaborate versions of the test. (It can also be noted that to our knowledge, no test that is valid has yet been developed for the setting that we use.)

[12] The cut-off point for "very strong" evidence – which is the highest category – is 10.

Table 2. Log marginal likelihoods – model with time-varying parameters and stochastic volatility.

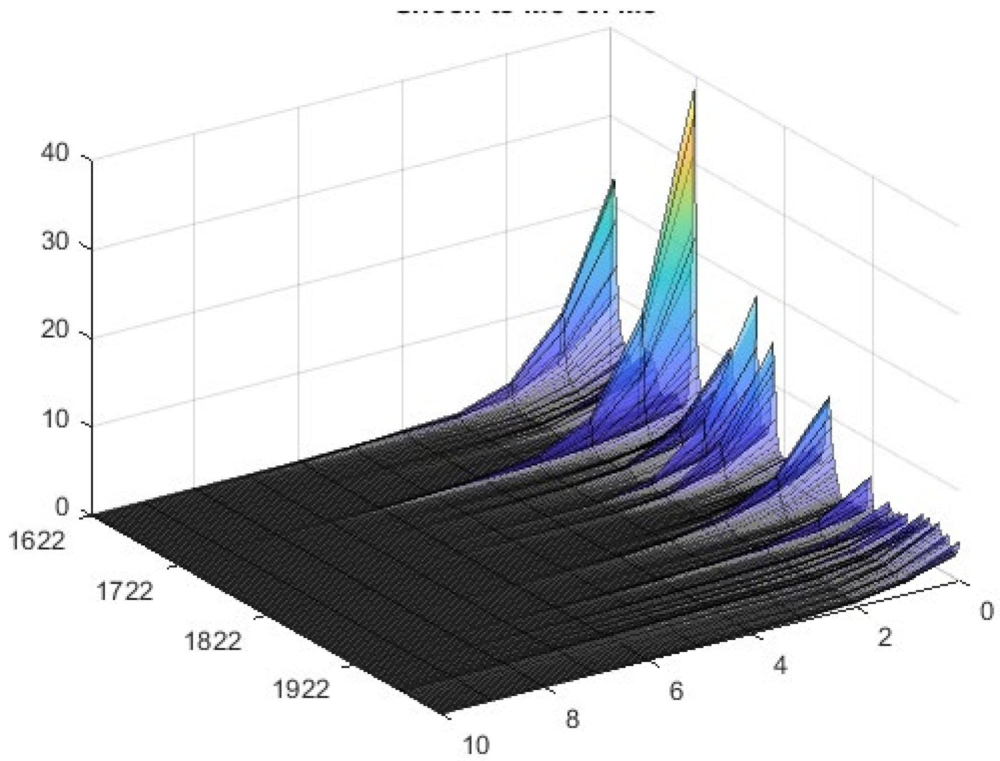
	Log marginal likelihood
Inflation endogenous	-2602.3
Inflation exogenous	-2611.1

Note: Model estimated on data ranging from 1622-2021.

Having established that the model in which inflation is endogenous with respect to money growth is preferred by the data, we next look at the properties of this model in more detail. First, we turn to the impulse-response functions. While our primary interest lies in the effect that a shock to money growth has on inflation, it is useful to first look at the effect that a shock to money growth has on money growth itself. This is shown in Figure 2. The size of the impulse is one standard deviation; the effect at horizon zero accordingly describes the estimated volatility of the shock to the money growth equation. Following this over time, it can be seen that there are some very substantial spikes. These spikes were related to a sudden expansion of money supply. Three of these are particularly worth pointing out. *i)* In the 1620s, copper coins were minted that contained less copper than their nominal value in anticipation of higher copper prices, while copper prices instead fell. If minting is conducted on government account – that is, there is no free minting – debased coins can continue to circulate at par with better coins but in the 1620s, the minting took place on a massive scale. Between 1624 and 1628 money supply trebled. *ii)* Around 1718, there was massive injection of coin tokens, which were the first fiat money that briefly came to dominate money supply. The money supply increased nearly five-fold between 1714 and 1718, the largest increase in such short period since 1620. *iii)* As a consequence of the suspension of the convertibility of Riksbank notes into silver coins in 1808-1809, at the time when Sweden was at war with Russia and lost Finland to the latter, money supply doubled.

Regarding the dynamic effect that shocks have had, there is some variation in how long-lived the effects have been. While the shocks were largest around 1718, they were most long lived in the 1970s.

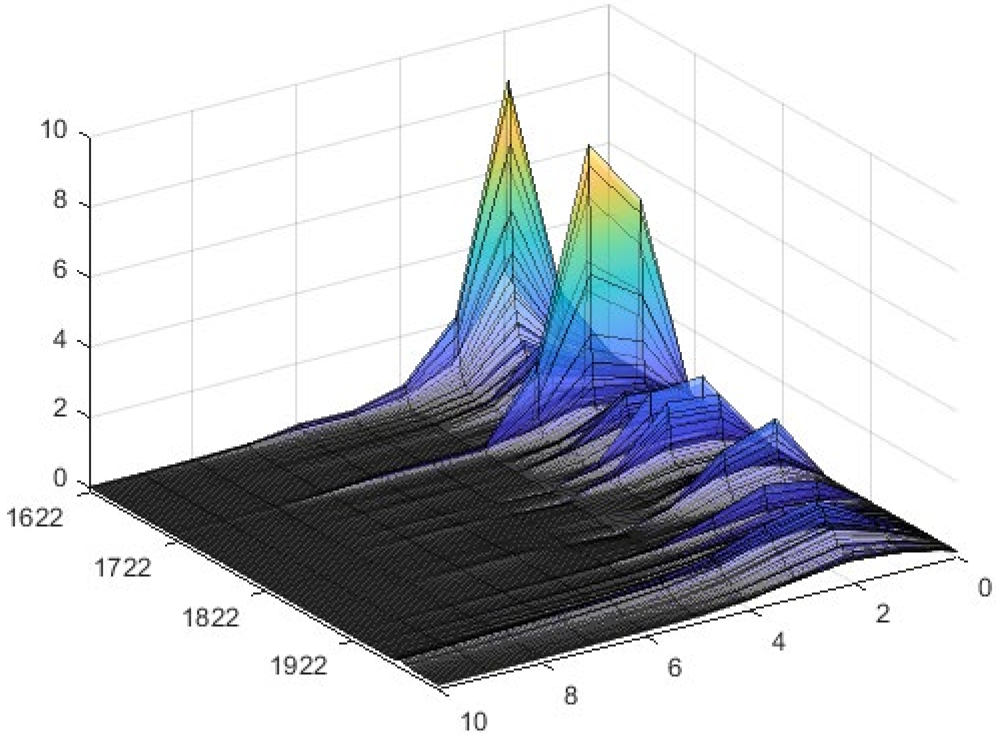
Figure 2. Impulse-response function: Effect of shocks to money growth on itself. Model with time-varying parameters and stochastic volatility.



Note: Size of impulse is one standard deviation. Effect in percentage points on vertical axis. Horizon in years and dates on horizontal axes.

Figure 3 displays the effect that a shock to money growth has had on inflation. As one would expect, inflation increases when money growth becomes unexpectedly high. Accordingly, we see some large effects on inflation when money growth shocks have been large, such as the episodes pointed out above (that is, the beginning of the sample, around 1718 and 1808-1809). But it is also clear from Figure 3 that the dynamic relation between the two variables has changed over time. Similar to what we saw concerning the effect of shocks to money growth on itself, there is variation in how long-lived the effects have been. For example, shocks to money growth were generally smaller in the 20th century than in the first half of the sample but their effect appears to have died out more slowly in the 20th century. One explanation for this is that in the pre-industrial society, short-term prices could increase dramatically in the short term due to harvest failures, but this was seen as supply shocks without long-lasting effects. Concerning the long-lasting effects of shocks in the 1970s, this is consistent with the problematic economic policies of that time. Wage formation functioned poorly in Sweden during this period and the economy was for many years subject to a price-wage spiral. Economic policy was focused on keeping the unemployment rate very low, rather than settling for the "natural" rate of unemployment. The credibility of Sweden's fixed exchange rate regime was very low, mainly because of the recurring devaluations of the krona – devaluations that contributed to the high (and persistent) inflation.

Figure 3. Impulse-response function: Effect of shocks to money growth on inflation. Model with time-varying parameters and stochastic volatility.



Note: Size of impulse is one standard deviation. Effect in percentage points on vertical axis. Horizon in years and dates on horizontal axes.

The final two impulse-response functions of the model – that is, the effect that shocks to inflation have on inflation itself and money growth respectively – are somewhat less interesting given the question posed in this paper. We nevertheless report them for completeness; they are shown in Figures A1 and A2 in the appendix. One feature worth pointing out here is the large variation that the volatility of shocks to inflation seem to have had. This again points to the need to use a model that allows for heteroskedasticity.

5.2 Out-of-sample results regarding Granger causality

In the previous section, we established that there was strong evidence that money growth Granger causes inflation. However, as we pointed out above, there is a large literature that favours the use of out-of-sample forecast performance as the relevant criterion when establishing whether there is Granger causality. In line with this, we next compare the two competing models from this perspective.

Our exercise is conducted the following way: We first estimate the two models on a sample ranging from 1622 to 1671.[13] Forecasts four years ahead (that is, covering the period 1672-1675) are generated and we record the forecast errors. We then expand the sample by one observation (that is, our sample is 1622 to 1672), re-estimate the models, generate new forecasts four years ahead (that is, covering the period 1673-1676) and again record the forecast errors. This continues until we run out of observations against which we can evaluate our forecasts. The last forecasts are accordingly based on the estimation sample 1622-2020.[14] This leaves us with a total of 350 forecast errors to evaluate for h=1, 349 for h=2, 348 for h=3 and 347 for h=4. The RMSFEs based on this exercise are shown in Table 3.

As can be seen from the top rows in the table, the model in which inflation is endogenous has a higher RMSFE than the model in which inflation is exogenous with respect to money growth regardless of which horizon we look at when the full dataset is used for evaluation. Differences between the two models are typically not large. The largest difference is found at the two-year horizon where the model in which inflation is endogenous has an RMSFE which is four percent higher. Two of the differences are found statistically significant based on the Diebold-Mariano test, namely at the three- and four-year horizons. Nevertheless, the RMSFEs indicate that money growth does not Granger cause inflation – a finding which stands in sharp contrast to the in-sample evidence based on marginal likelihoods presented in Section 5.1.

The finding that the model in which inflation is endogenous does not improve the forecasts relative to the model in which inflation is exogenous with respect to money growth might seem somewhat surprising given our results in the in-sample analysis. Aiming to try to understand why we get this result, we next break down our results according to the sub-samples that we identified in Section 3. We accordingly have a closer look at the RMSFEs for the periods 1672-1776, 1777-1873, 1874-1971 and 1972-2020; these are also presented in Table 3.

13 While we want as many forecasts as possible for our evaluation, one should also use enough observations for the estimation so that the model can be estimated with a fair amount of precision. We believe that our choice here strikes a reasonable trade-off between these two aspects.

14 Since 2021 is the last observation available for comparing forecasts to outcomes, only the one-year-ahead forecast can be evaluated for the last out-of-sample forecast.

Table 3. RMSFEs for different periods from out-of-sample forecast exercise.

Period	Forecast horizon (years)	Inflation endogenous	Inflation Exogenous
1672-2021	1	11.14	11.07
	2	11.59	11.15
	3	11.33	11.12[b]
	4	11.04	10.94[a]
1672-1776	1	18.15	18.00
	2	18.62	17.63
	3	17.92	17.51[b]
	4	17.22	17.05[b]
1777-1873	1	7.30	7.06
	2	7.18	7.11
	3	7.15	7.15
	4	7.13	7.16
1874-1971	1	5.99	6.25
	2	7.25[a]	7.44
	3	7.56	7.43
	4	7.62	7.48[a]
1972-2021	1	2.00	2.15
	2	2.57	2.84
	3	2.93	3.13
	4	3.25	3.31

Note: a, b and c indicate significance of Diebold-Mariano test at the 1, 5 and 10 percent significance level.

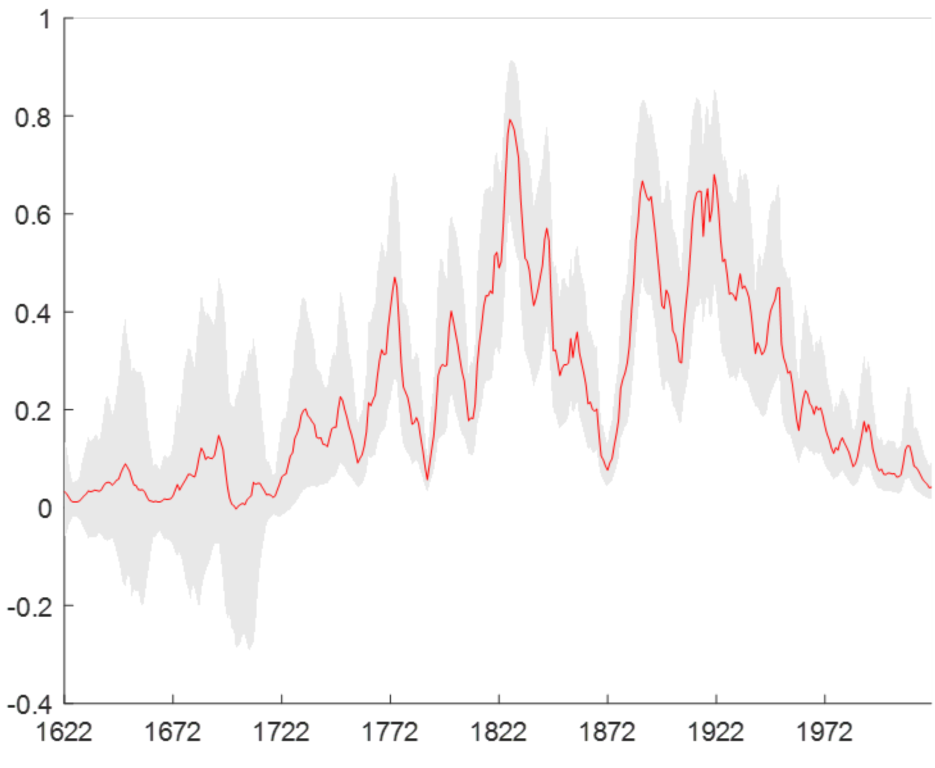
Looking at the results from the four sub-periods, we note that the evidence is more mixed than when looking at the full period. The model in which inflation is endogenous has a lower RMSFE at the four-year horizon for the 1777-1873 period and at the one- and two-year horizons for the 1874-1971 period; in addition, its RMSFEs are lower at all four horizons for the 1972-2021 period. Again though, differences are typically small; of the 16 pairs of RMSFEs for the sub-periods, the Diebold-Mariano test only finds four of them statistically significant – three of which are in favour of the model in which inflation is exogenous with respect to money growth. Only at the two-year horizon for the 1874-1971 period do we find significant evidence in favour of Granger causality.

Summing up our results from the out-of-sample forecast exercise, it appears that the relative predictive ability varies somewhat over time. It is also clear that the out-of-sample analysis does not support our finding from the within-sample analysis that money growth Granger causes inflation. Such contradictions are unfortunately not uncommon in the literature and a likely explanation is the well-known issue of in-sample overfitting –

that is, that a more complex model often finds support in the data within-sample but that this added complexity does not pay off when forecasting.

Finally, we want to point out that the lack of evidence in favour of Granger causality that we have established in this sub-section does not mean that money growth is irrelevant for inflation. Looking at Granger causality, we have focused on the *dynamic* relationship between the two variables. However, there is also a *contemporaneous* relation between inflation and money growth. In the impulse-response functions from the model, this relation is reflected in the contemporaneous effect that shocks to inflation had on money growth (but not vice versa); this is shown in Figure A2 in the appendix. The reason for this is the recursive structure of the structural model employed and the fact that we chose to define the vector of dependent variables as $\boldsymbol{y}_t = (\pi_t \quad \mu_t)'$. However, as was pointed out above (in Section 4.1), we believe that in practice it is reasonable to assume that both variables can affect each other within the year. In order to illustrate this issue, Figure 4 shows the estimated time-varying correlation between the reduced-form disturbances of the model (see footnote 5). This is calculated as $\rho_t = \widehat{\omega}_{12t}/\left(\sqrt{\widehat{\omega}_{11t}}\sqrt{\widehat{\omega}_{22t}}\right)$ where $\widehat{\omega}_{12t}$, $\widehat{\omega}_{11t}$ and $\widehat{\omega}_{22t}$ are the relevant elements from the estimated reduced-form covariance matrix $\widehat{\boldsymbol{\Omega}}_t = \widehat{\boldsymbol{B}}_{0t}^{-1}\widehat{\boldsymbol{\Sigma}}_t\left(\widehat{\boldsymbol{B}}_{0t}^{-1}\right)'$.

Figure 4. Estimated correlation between reduced-form disturbances to inflation and money growth.



Note: Grey bands are 68 percent credible intervals.

As can be seen from Figure 4, there is typically a positive correlation between the disturbances. The strength of the correlation varies over time. From 1622 to the mid-1700s and near the end of the sample, the correlation is low – typically less than 0.2. Peaks, that reach above 0.6, are found around the 1820s, before convertibility of banknotes into silver coins was restored in 1834 (Lobell, 2010, p. 298), and around the first world war, when the gold standard was suspended. Obviously, we cannot draw any conclusions regarding causality here seeing that we are looking at correlations from a reduced form. However, we note that our findings are consistent with money growth generating inflation within the year, such as the episode of 1718 mentioned above; it is, of course, also consistent with inflation generating money growth. Taken together, we conclude that the results presented in this sub-section indicate that money growth likely matters for inflation but that it does not do so dynamically out-of-sample.

6. Conclusions

In this paper we have assessed whether money growth is Granger causal for inflation using a historical dataset. The within-sample analysis strongly suggests that money supply has predictive power for inflation so that monetary shocks – that is, a higher growth rate of the money supply than expected – are followed by higher inflation. Even if Granger causality may not involve actual causality, actual causality can be demonstrated in several cases by historical studies, for example, how inflation directly followed – often within months – extreme expansions of fiat money (Jörberg 1972; Edvinsson and Söderberg, 2010). Still, we cannot rule out that some of the covariations are due to randomness.

The within-sample analysis also indicates that monetary shocks were more volatile during the pre-industrial period. We also find that these shocks seem to have been longer-lived during parts of the 20th century than in the pre-industrial period – a finding which reflects differences in the structure of the Swedish economy and likely also how inflation expectations were formed. Pre-industrial supply shocks – often stemming from the agricultural sector – did not have long-lasting effects whereas shocks that occurred in the post-Bretton Woods era (but before inflation targeting was adopted) hit an economy that was subject to a price-wage spiral.

In contrast, the out-of-sample analysis conducted indicates that there is little evidence of money growth Granger causing inflation. In fact, looking at the full sample, statistically significant differences against Granger causality can be found at the two longest forecast horizons; statistically significant support in favour of Granger causality can only be found at the two-year horizon for the 1874-1971 period. It appears that the relative predictive ability that money growth has on inflation varies somewhat over time. This seems reasonable and we note that it is in line with a conclusion from a literature focused more on modern macroeconomic aspects of the Granger causality of money growth for inflation, namely that the regime likely matters (e.g. Berger and Österholm, 2011a, 2011b; Borio *et al.*, 2023).

Concerning our empirical results, we also note that even if the out-of-sample evidence found little evidence in favour of Granger causality, this does not mean that money growth does not matter for inflation. Our data indicate that there tends to be a positive *contemporaneous* relation between money growth and inflation. This means that even if there are no dynamic effects of money growth on inflation, one cannot exclude that money growth affects inflation.

Finally, we want to say a few words about directions for future research. Annual data are admittedly a bit crude and, as mentioned above, they mean that the recursive structure of the model can be questioned. With quarterly data, such an assumption can provide a better (though still not perfect) approximation to reality. It would accordingly be desirable to collect historical data at higher frequency, for example, quarterly data for several centuries. At least for Sweden, there are empirical material in the archives that would allow such an endeavour, although it would be very time consuming. Access to such data would, however, be appreciated by both economic historians and macroeconomists.

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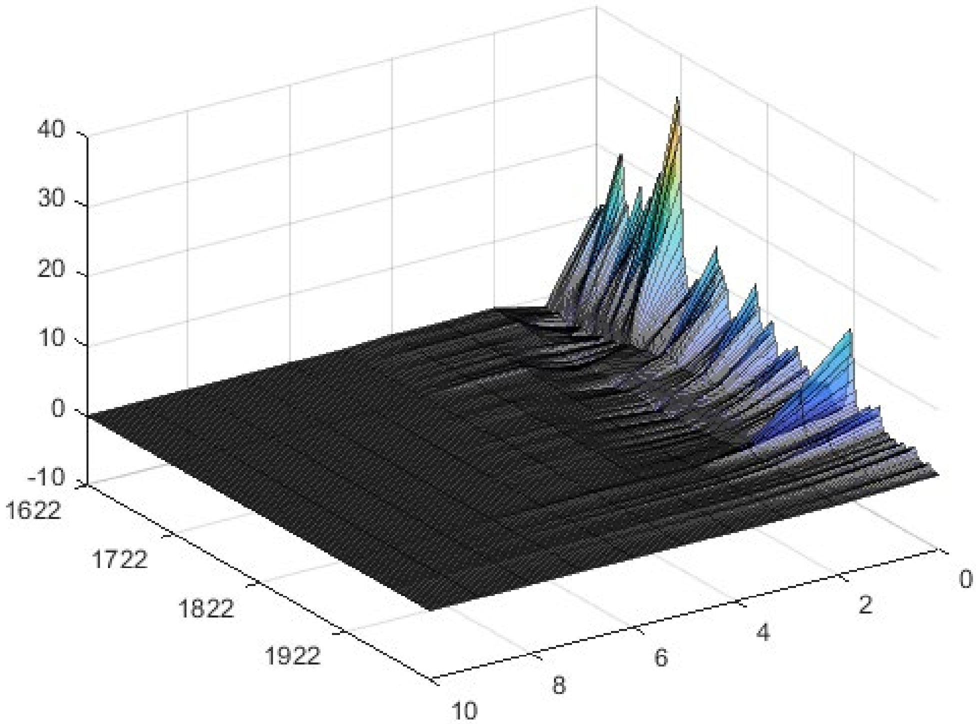
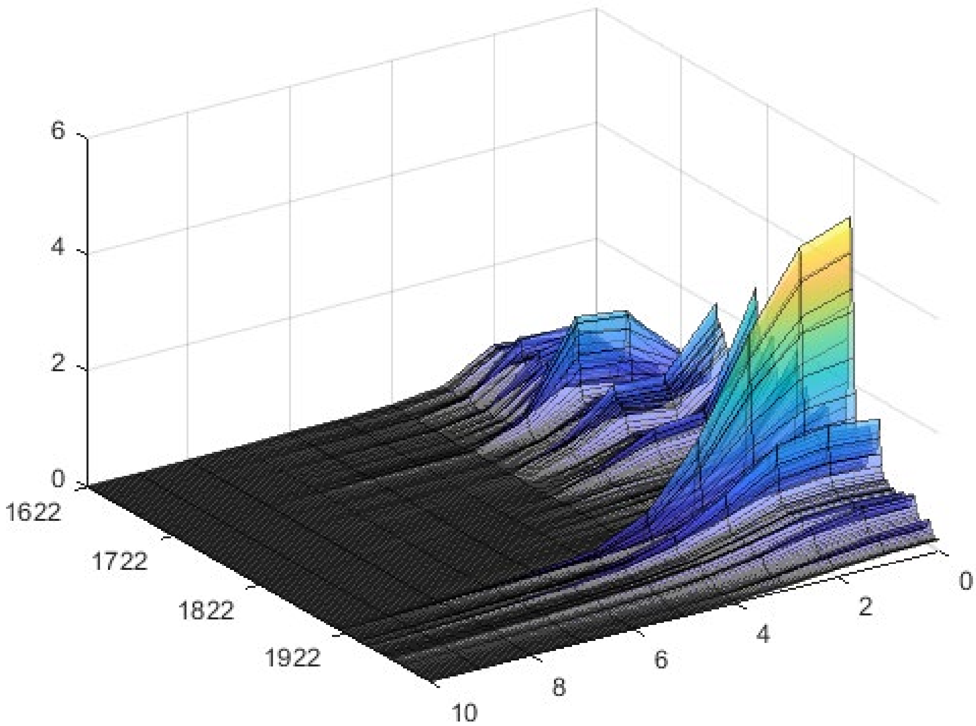
Appendix

Figure A1. Impulse-response function: Effect of shocks to inflation on itself. Model with time-varying parameters and stochastic volatility.



Note: Size of impulse is one standard deviation. Effect in percentage points on vertical axis. Horizon in years and dates on horizontal axes.

Figure A2. Impulse-response function: Effect of shocks to inflation on money growth. Model with time-varying parameters and stochastic volatility.



Note: Size of impulse is one standard deviation. Effect in percentage points on vertical axis. Horizon in years and dates on horizontal axes.