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Does Inflation Cause Gold Market Price Changes? Evidence on the G7 Countries from the Tests of Nonparametric Quantile Causality in Mean and Variance

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Abstract

This paper utilises the newly proposed nonparametric causality-in-quantiles test to examine the predictability of mean and variance of changes in gold prices based on inflation for G7 countries. The causality-in-quantiles approach permits us to test for not only causality in mean but also causality in variance. We start our investigation by utilising tests for nonlinearity. These tests identify nonlinearity, showing that the linear Granger causality tests are subject to misspecification error. Unlike tests of misspecified linear models, our nonparametric causality-in-quantiles tests find causality in mean and variance from inflation to gold market price changes between the 0.20 quantile and the 0.70 quantile, implying that very low- and high- price changes in gold markets are not related to inflation. These changes should be related to other sources, such as financial shocks and exchange market shocks. We find support that gold serves as a hedge against inflation, but only in the mid-quantile ranges, i.e., quantiles from 0.20 to 0.70. Our results show that gold does not serve as a hedge against inflation during periods when gold market price changes are very low or very high, which are respectively quiet and highly volatile periods.

Keywords: Gold, Inflation, Spot and futures markets; Quantile causality.

JEL Codes: C22, Q02, E31.

I. Introduction

Gold, the oldest medium of exchange, is a popular hedge (Chang et al. 2013), has symbolic value (Godness et al. 2016), is a broad measure of economic conditions (Natanelov et al. 2011) and is the only true standard of value, even in war and crisis times. It has a great impact on economies and is commonly considered to be a safe haven for investment. Gold is one of the rarest elements, representing just 0.0000001% or one part per billion. Its unique properties make gold irreplaceable for all practical purposes. Gold is traded in one of the largest global spot and futures market. Broadly, gold market refers to the continuous trading of gold throughout the world based on the intra-day spot price, derived from over-the-counter (OTC) trading across the world. A large futures trading also takes place in the gold market. Due to its unique properties, it is the most popular precious metals as an investment, making the gold market one of the largest global OTC markets. Gold is generally bought for diversifying risk, especially through the use of futures contracts and derivatives. The financial market for gold is quite robust, which allows market participants to use gold as a store of wealth, an investment, and as a source of high quality collateral.

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As it holds its value independently of domestic politics, its stability against fluctuating currencies and people's high interest and confidence over the centuries, the global gold market and its interdependency with other commodities and economic dynamics have always attracted a great deal of economic curiosity. Moreover, due to its high profit potential and remarkable risk protection feature, the gold market is receiving increasing attention (Zhang and Wei 2010). Therefore, consistent interest in the gold market has necessitated inquiry regarding its relationship with other macro-economic factors, such as oil and inflation.

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Although there is a large body of research that either supports the role of gold as a traditional hedge against inflation (Jaffe 1989, Dempster and Artigas 2009, Narayan et al. 2010, Ghosh 2011) or not (Soytas et al. 2009) and as a hedge in bond and stock markets (Baur and Lucey 2010), relatively few studies have investigated the causal relationship between inflation and gold prices. To our knowledge, no study has employed the nonparametric quantile-based test that we use in this study. The nonparametric causality-in-quantiles test we use is robust to nonlinearities, structural breaks and outliers.

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Despite the limited body of literature regarding the effect of inflation on gold market price changes, several studies have analysed the impact of gold prices on predicting inflation (Mahdavi and Zhou 1997), the hedging effectiveness of gold and its ability to forecast future inflation (Godness et al. 2016, Beckmann and Czudaj 2013, Wang et al. 2011), the long-term relationship between the price of gold and inflation (Ghosh et al. 2004, Worthington and Pahlavani 2007), the accuracy of gold as an indicator of future inflation over other measures (such as consumer price index and oil), its effectiveness in gauging and combating the effects of inflation on a portfolio (Ranson 2005), the ineffectiveness of unexpected changes in the CPI in relation to gold prices (Blose 2010) or the use of inflation to examine the dependence structure and linkages between gold and other commodity markets, especially oil (Zhang and Wei 2010, Narayan et al. 2010, Reboredo 2013, Tiwari and Sahadudheen 2015).

□

Evidence on the effect of the inflation on gold market price changes from available studies is mixed, which could be due to the different country-specific case studies, data samples and methodologies, and misspecification errors. However, no insightful work has been done yet to examine the dynamic link between inflation and gold market price changes that considers data issues, such as the structural breaks, outliers, and nonlinearity. Against this backdrop, the purpose of this study is to examine the Granger causality in mean and variance from inflation to gold market price changes in G7 countries using the nonparametric causality-in-quantiles test. This study investigates the relationship between changes in gold prices and inflation using a novel methodology (Balciar et al. 2016a, b) that is useful in detecting nonlinear causality.

□

To address the gap in the literature about the dynamic links between inflation on gold markets, we utilise the novel nonparametric causality-in-quantiles test of Balciar et al. (2016a, b) to study the predictability of the conditional mean and variance of gold market price changes based on inflation. The nonparametric causality-in-quantiles test combines the test for the nonparametric causality of the k -th order by Nishiyama et al. (2011) with the causality-in-quantiles test by Jeong et al. (2012) and, hence, can be thought of as a generalization of the former. The causality-in-quantiles approach used in this study is novel in three ways. First, it is robust to misspecification errors, as it detects the underlying linear and nonlinear dependence structure in the time series under examination. Second, using this approach, one can test not only causality in mean (1st moment) but also causality in variance

(2nd moment). Thus, higher-order dependency can be studied. Such an examination is vital in light of the fact that, during some periods, causality in the conditional mean may not exist, whereas higher-order interdependencies may be significant. This method enables us to test the mean and variance causality links from inflation to gold price changes.

□

In this paper, we use monthly data from the December 1979-August 2016 period. To the best of our knowledge, this is the first study to analyse the predictability of mean and the volatility of changes in gold prices based on inflation using the nonparametric causality-in-quantiles method. The paper obtains new findings on the effect of inflation on gold market price changes. Causality-in-quantiles tests used in this study usually find causality from inflation to gold market price changes between the 0.20 quantile and the 0.70 quantile, meaning that very low- and high-price movements in gold markets are not related to inflation. These results should be connected to different factors such as financial shocks and exchange market shocks. The empirical evidence shows that gold serves as a hedge against inflation, a commonly known fact. The causality is not only in the mean (price changes) but also in the variance, meaning that inflation impacts both gold market price changes and gold market volatility.

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The rest of the paper is organized in the following manner. Section 2 explains the methodology. Section 3 presents the data and empirical evidence. Finally, Section 4 concludes the paper.

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2. Methodology

We utilise the nonparametric causality-in-quantiles approach of Balcilar et al. (2016a, b), a novel method that expands upon the approaches of Nishiyama et al. (2011) and Jeong et al. (2012). This methodology is capable of accounting for nonlinear causality in quantiles through a hybrid approach. The price changes on gold are denoted as y_t , and inflation is denoted as x_t . As described in Jeong et al. (2012), we characterize quantile-based causality as follows¹: x_t does not cause y_t in the θ -th quantile with respect to the lag vector of $\{y_{t-1}, \dots, y_{t-p}, x_{t-1}, \dots, x_{t-p}\}$ if

□

¹ The explanation in this section closely follows those in Nishiyama et al. (2011) and Jeong et al. (2012).

$$Q_\theta(y_t | y_{t-1}, \dots, y_{t-p}, x_{t-1}, \dots, x_{t-p}) = Q_\theta(y_t | y_{t-1}, \dots, y_{t-p}) \quad (1)$$

x_t is presumably the cause of y_t in the θ -th quantile with regard to $\{y_{t-1}, \dots, y_{t-p}, x_{t-1}, \dots, x_{t-p}\}$ if

$$Q_\theta(y_t | y_{t-1}, \dots, y_{t-p}, x_{t-1}, \dots, x_{t-p}) \neq Q_\theta(y_t | y_{t-1}, \dots, y_{t-p}) \quad (2)$$

Here, $Q_\theta(y_t | \cdot)$ is the θ -th quantile of y_t . The conditional quantiles of y_t , $Q_\theta(y_t | \cdot)$, depends on t , and the quantiles are constrained between zero and one, i.e., $0 < \theta < 1$.

For a compact presentation of the causality-in-quantiles tests, we define the following vectors: $Y_{t-1} \equiv (y_{t-1}, \dots, y_{t-p})$, $X_{t-1} \equiv (x_{t-1}, \dots, x_{t-p})$, and $Z_t = (X_t, Y_t)$. Additionally, we define the conditional distribution functions $F_{y_t|Z_{t-1}}(y_t | Z_{t-1})$ and $F_{y_t|Y_{t-1}}(y_t | Y_{t-1})$, with the distribution functions of y_t conditioned on vectors Z_{t-1} and Y_{t-1} , respectively. Moreover, the conditional distribution function $F_{y_t|Z_{t-1}}(y_t | Z_{t-1})$ is assumed to be continuous in y_t for almost all Z_{t-1} . By defining $Q_\theta(Z_{t-1}) \equiv Q_\theta(y_t | Z_{t-1})$ and $Q_\theta(Y_{t-1}) \equiv Q_\theta(y_t | Y_{t-1})$, we can see that $F_{y_t|Z_{t-1}}\{Q_\theta(Z_{t-1}) | Z_{t-1}\} = \theta$, which holds true with a probability equal to one. Subsequently, the hypotheses to be tested for causality in quantiles, taking into account conditions (1) and (2), can be stated as:

$$H_0: P\{F_{y_t|Z_{t-1}}\{Q_\theta(Y_{t-1}) | Z_{t-1}\} = \theta\} = 1 \quad (3)$$

$$H_1: P\{F_{y_t|Z_{t-1}}\{Q_\theta(Y_{t-1}) | Z_{t-1}\} = \theta\} < 1 \quad (4)$$

To define a measurable metric for the practical implementation of the causality-in-quantiles tests, Jeong et al. (2012) use the distance measure $J = \{\varepsilon_t E(\varepsilon_t | Z_{t-1}) f_Z(Z_{t-1})\}$, where ε_t denotes the regression error and $f_Z(Z_{t-1})$ denotes the marginal density function of Z_{t-1} . Thus, the causality-in-quantiles test depends on the regression error ε_t . The regression error ε_t arises based on the null hypothesis specified in equation (3), which holds true if and only if $E[\mathbf{1}\{y_t \leq Q_\theta(Y_{t-1}) | Z_{t-1}\}] = \theta$. To make the regression error explicit, we rewrite this last statement as $\mathbf{1}\{y_t < Q_\theta(Y_{t-1})\} = \theta + \varepsilon_t$, where $\mathbf{1}\{\cdot\}$ is an indicator function. Following Jeong et al. (2012) and based on the regression error, the distance metric can be characterized as:

$$J = E \left[\left\{ F_{y_t|Z_{t-1}} \{ Q_\theta(Y_{t-1}) | Z_{t-1} \} - \theta \right\}^2 f_Z(Z_{t-1}) \right] \quad (5)$$

Under conditions (3) and (4), the distance metric in (5) satisfies $J \geq 0$. The statement holds true with an equality, i.e., $J = 0$, if and only if the null hypothesis H_0 in equation (3) is true, while $J > 0$ holds true under the alternative hypothesis H_1 in equation (4). The feasible counterpart of the distance measure J in equation (5) gives us a kernel-based causality-in-quantiles test for the fixed quantile θ and is characterized as:

$$\hat{J}_T = \frac{1}{T(T-1)h^{2p}} \sum_{t=p+1}^T \sum_{s=p+1, s \neq t}^T K \left(\frac{Z_{t-1} - Z_{s-1}}{h} \right) \hat{\varepsilon}_t \hat{\varepsilon}_s \quad (6)$$

where $K(\cdot)$ represents a known kernel function, h is the bandwidth for the kernel estimation, T indicates the sample size, and p denotes the lag order used to define vector Z_t . Jeong et al. (2012) show that the re-scaled statistic $T h^p \hat{J}_T / \hat{\sigma}_0$ is asymptotically distributed as a standard normal distribution, where $\hat{\sigma}_0 = \sqrt{2\theta(1-\theta)} \sqrt{1/(T(T-1)h^{2p})} \sqrt{\sum_{t \neq s} K^2((Z_{t-1} - Z_{s-1})/h)}$. The most critical component of the test statistic \hat{J}_T is the regression error $\hat{\varepsilon}_t$. In our specific case, the estimator of the unknown regression error is characterized as:

$$\hat{\varepsilon}_t = \mathbf{1}\{y_t \leq \hat{Q}_\theta(Y_{t-1})\} - \theta \quad (7)$$

In equation (7), the quantile estimator $\hat{Q}_\theta(Y_{t-1})$ yields an estimate of the θ -th conditional quantile of y_t given Y_{t-1} . We evaluate $\hat{Q}_\theta(Y_{t-1})$ by utilising the nonparametric kernel approach:

$$\hat{Q}_\theta(Y_{t-1}) = \hat{F}_{y_t|Y_{t-1}}^{-1}(\theta|Y_{t-1}) \quad (8)$$

where $\hat{F}_{y_t|Y_{t-1}}(y_t|Y_{t-1})$ represents the Nadarya-Watson kernel estimator given by:

$$\hat{F}_{y_t|Y_{t-1}}(y_t|Y_{t-1}) = \frac{\sum_{s=p+1, s \neq t}^T L\left(\frac{Y_{t-1} - Y_{s-1}}{h}\right) \mathbf{1}\{y_s \leq y_t\}}{\sum_{s=p+1, s \neq t}^T L\left(\frac{Y_{t-1} - Y_{s-1}}{h}\right)} \quad (9)$$

In this equation, $L(\cdot)$ signifies a known kernel function and h is the bandwidth utilised in the kernel estimation.

The causality in variance suggests an impact on the volatility, which may exist even when there is no causality in the mean (1st moment). Testing for Granger causality in second or higher moments has some complications, and the procedure for such tests should be carefully defined since rejection of causality in the moment m does not imply non-causality in the moment k for $m < k$. We begin by employing Nishiyama et al.'s (2011) nonparametric Granger quantile causality method. To demonstrate the causality in higher-order moments, we first consider the process:

$$y_t = g(Y_{t-1}) + \sigma(X_{t-1})\varepsilon_t \quad (10)$$

where ε_t denotes an independently and identically distributed (i.i.d.) process and the unknown functions $\sigma(\cdot)$ and $g(\cdot)$ satisfy some properties that are sufficient for the stationarity of y_t . Although this representation does not permit linear or non-linear causalities from X_{t-1} to y_t , it does allow X_{t-1} to have predictive power for y_t^2 when $\sigma(\cdot)$ is an established nonlinear function. The representation in equation (10) illustrates that squares for X_{t-1} are not necessarily entered into the nonlinear function $\sigma(\cdot)$. Therefore, we re-specify equations (3) and (4) into a null hypothesis H_0 and an alternative hypothesis H_1 for causality in variance as follows:

$$H_0: P\left\{F_{y_t^2|Z_{t-1}}\{Q_\theta(Y_{t-1})|Z_{t-1}\} = \theta\right\} = 1 \quad (11)$$

$$H_1: P\left\{F_{y_t^2|Z_{t-1}}\{Q_\theta(Y_{t-1})|Z_{t-1}\} = \theta\right\} < 1 \quad (12)$$

To obtain a feasible test statistic to test the null hypothesis H_0 in equation (11), we substitute y_t in equations (6) to (9) with y_t^2 . A problem may arise with the causality test based on the definition given in equation (10), as there may be causality in the second moment (variance) along with causality in the conditional first moment (mean). We can illustrate this with the following model:

$$y_t = g(X_{t-1}, Y_{t-1}) + \varepsilon_t \quad (13)$$

Thus, higher-order causality in quantiles can be stated as:

$$H_0: P\{F_{y_t^k|Z_{t-1}}\{Q_\theta(Y_{t-1})|Z_{t-1}\} = \theta\} = 1 \quad \text{for } k = 1, 2, \dots, K \quad (14)$$

$$H_1: P\{F_{y_t^k|Z_{t-1}}\{Q_\theta(Y_{t-1})|Z_{t-1}\} = \theta\} < 1 \quad \text{for } k = 1, 2, \dots, K \quad (15)$$

Fully incorporating the concepts, we specify that x_t Granger-causes y_t in quantile θ up to the K -th moment utilising equation (14) to formulate the test statistic of equation (6) for each k . Nishiyama et al. (2011) construct nonparametric Granger causality tests using the density-weighted approach as in Jeong et al. (2012) and show that density-weighted nonparametric tests in higher moments have the same asymptotic normal distribution as the test for causality in the first moment, although some stronger moment conditions might be necessary. Nevertheless, it is not an easy task to test for all $k = 1, 2, \dots, K$ jointly, as the statistics are correlated (Nishiyama et al., 2011). To overcome this issue, one can follow the sequential testing approach in Nishiyama et al. (2011) to test for causality in both models defined in equations (10) and (13). In this approach, we first test for nonparametric Granger causality in the first moment ($k = 1$) but still continue to test for causality in variance even if non-causality is not rejected. That is, if the null hypothesis for $k = 1$ is not rejected, then there might still be causality in the second moment, and thus, we conduct the tests for $k = 2$. This methodology permits us to test for the presence of causality in variance only as well as causality in the mean and the variance successively. That is, we can examine the existence of causality in mean and causality in variance sequentially. The empirical analysis of nonparametric causality in quantiles requires three specifications: the lag order p , the bandwidth h , and the kernel type for $K(\cdot)$ and $I_r(\cdot)$ in equations (6) and (9), respectively. In this study, we use the lag order of 1 based on the Schwarz information criterion (SIC) under a VAR involving inflation and gold market price changes. Additionally, in regard to choosing lags, the SIC is considered to be more parsimonious than other lag-length selection criteria. The SIC helps to overcome the issue of over-parameterization which usually arises in

nonparametric frameworks.² The bandwidth value is chosen by employing least-squares cross-validation techniques.³ Finally, for $K(\cdot)$ and $L(\cdot)$, Gaussian-type kernels are used.

I

3. Data and Empirical Findings

In the empirical analysis, we use the monthly US dollar⁴ prices of gold spot and futures contracts traded on the London Bullion Market (LBMA) and the seasonally adjusted Consumer Price Index (CPI) as proxies for price level. Note that the gold future variables represent the gold futures markets at twelve maturities (1-month, 2-month, 3-month, 4-month, 5-month, 6-month, 7-month, 8-month, 9-month, 10-month, 11-month and 12-month). The CPI data is used in monthly frequencies for the period of December 1979-August 2016 period. The CPI series are taken from International Financial Statistics (IFS). Inflation is measured by the monthly difference in the log CPI [$100 \times \log(CPI_t/CPI_{t-1})$]. Inflation series are stationary, while gold spot and futures price series are non-stationary in log levels, as indicated by standard unit root tests⁵. Since the methodology requires stationary data, we use gold market price change series, which are obtained as the first differences of the natural logarithmic values of the gold price series expressed as percentages. Figure 1 shows the time series plot of the inflation series of G7 countries, while Figure 2 plots the log changes (%) of the spot and futures of the gold price series.

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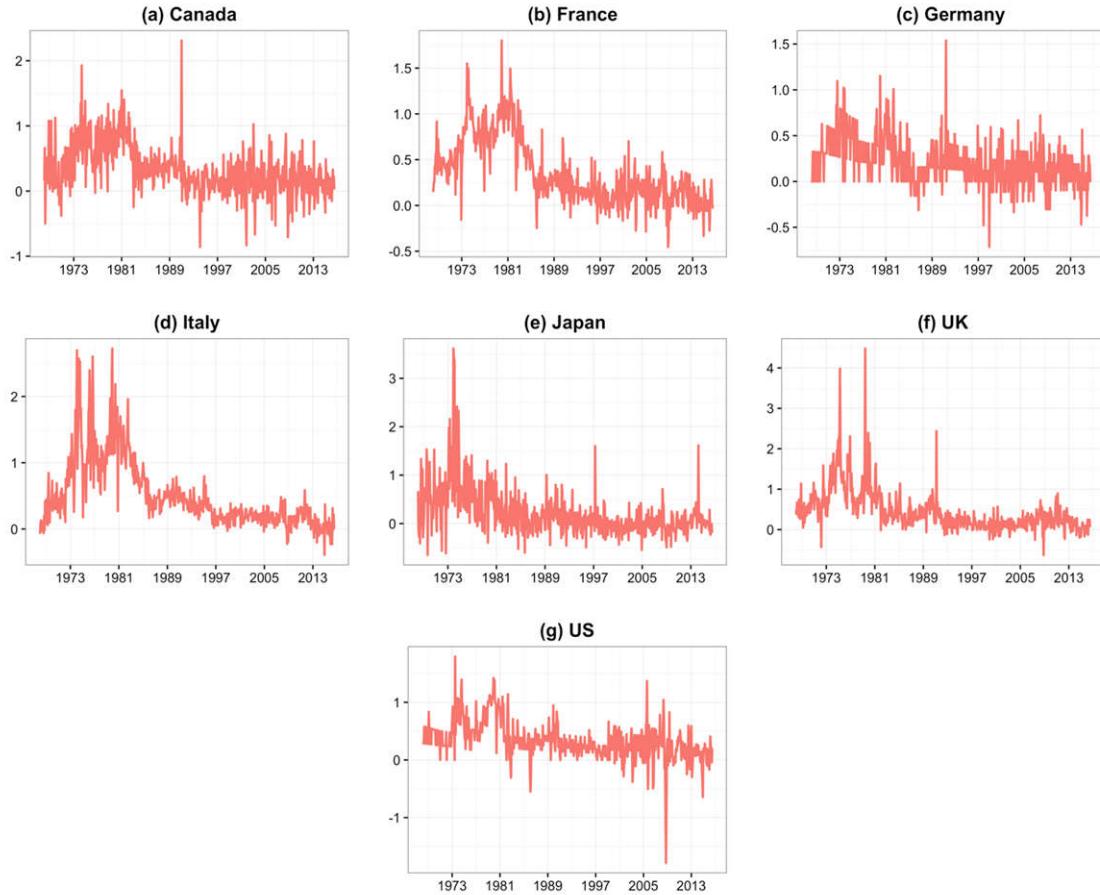
We begin with a brief discussion of selected key features of the series. In Table 1, we report the mean, standard deviation, kurtosis, skewness, Jarque-Bera normality test (JB), Ljung-Box first [Q(1)] and fourth [Q(4)] autocorrelation tests, and first- [ARCH(1)] and fourth-order [ARCH(4)] Lagrange multiplier (LM) tests for autoregressive conditional heteroskedasticity (ARCH) for the inflation series of G7 countries and gold spot and futures contracts. The mean of the inflation series suggests that it is lowest for Germany but gradually increases for Italy.

² Hurvich and Tsai (1989) examine the Akaike Information Criterion (AIC) and show that it is biased towards selecting an over-parameterized model, while the SIC is asymptotically consistent.

³ For each quantile, we determine the bandwidth h using the leave-one-out least-squares cross-validation method of Racine and Li (2004) and Li and Racine (2004).

⁴ We additionally performed all tests after converting spot and 1- to 12-month futures gold prices into national currencies, i.e., yen for Japan, sterling for UK, Canadian dollar for Canada, and euro for France, Germany, and Italy. The causality-in-quantiles test results obtained using gold prices in local currency units were qualitatively the same with the results presented in the paper, the causality from inflation to gold market price changes were even stronger for some countries. These results are not reported to save space, but available from the authors upon request.

⁵ Complete details of the unit root tests are available from the authors upon request.

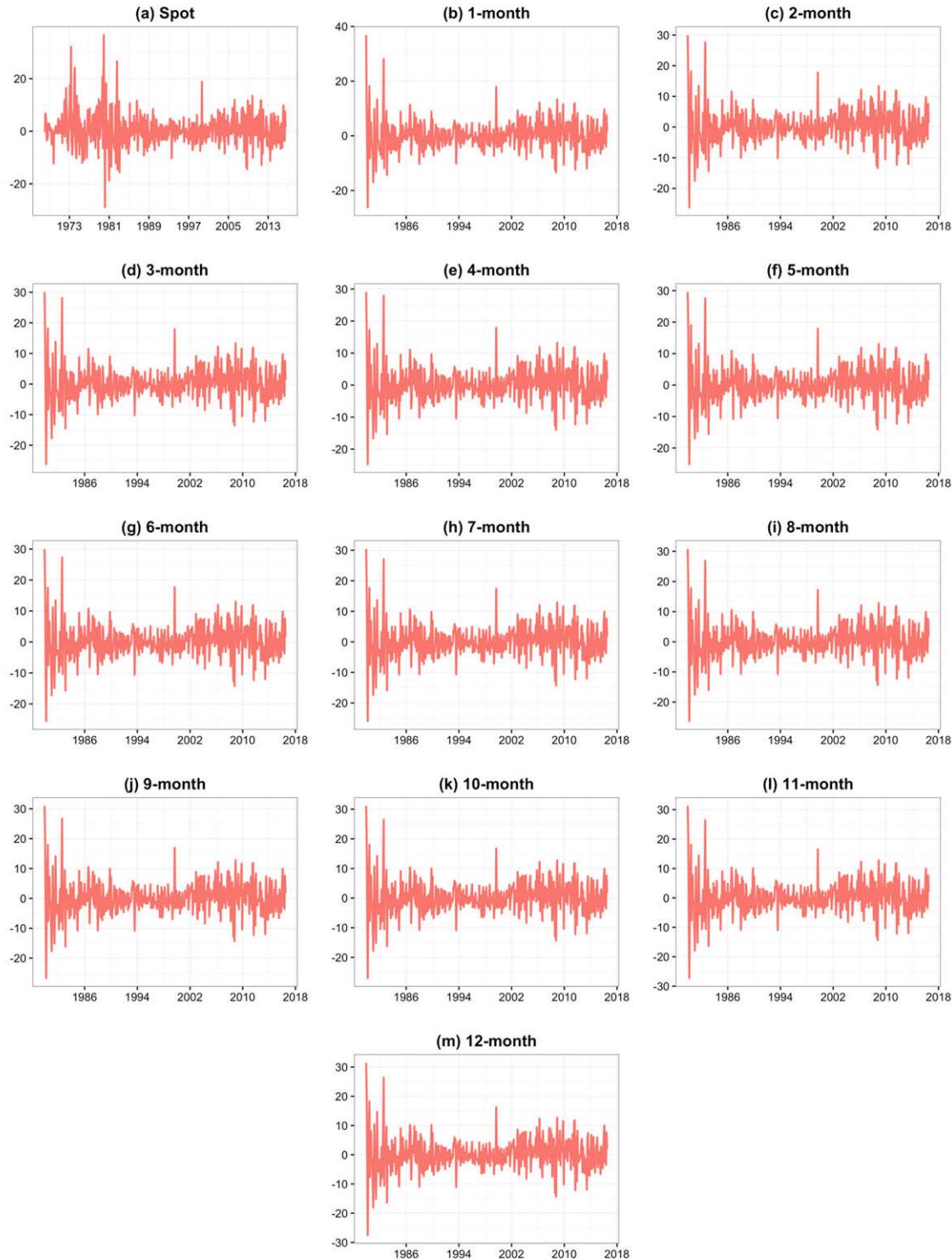


Note: The figure plots monthly inflation rates in percentages.

Figure 1. Time series plots of inflation rates

The mean of the gold market price changes suggests that it is approximately 0.24. Second, we observe that the inflation series are less volatile than changes in gold market prices. The positive values of the skewness statistic suggest a smaller probability of increases in inflation and gold market price changes. All series distributions have fat tails, in accordance with the high values of the kurtosis statistics. More importantly, for our context of causality-in-quantiles, both variables are skewed to the right with excess kurtosis, resulting in non-normal distributions, as indicated by the strong rejection of the Jarque-Bera statistic at a 1% significance level. The heavy tails of the distributions in both series provide preliminary justification for the nonparametric causality-in-quantiles test used in this paper. The Ljung-Box tests indicate the existence of a first-order serial correlation in the inflation series of G/ countries, while there is no first-order serial correlation in the gold market price changes. However, the Ljung-Box tests do not refute the nonexistence of a fifth-order serial correlation in the inflation series and the gold market price changes. The autoregressive conditional

heteroskedasticity-Lagrange multiplier (ARCH-LM) statistic suggests that ARCH effects exist in all series.



Note: The figure plots the log price changes (%) of the gold markets.

Figure 2. Time series plots of spot and futures gold market price changes

Table 1. Descriptive statistics for inflation and gold market price changes (%)

	<i>n</i>	Mean	S.D.	Min	Max	Skewness	Kurtosis	JB	O(1)	O(4)	ARCH(1)	ARCH(4)	Sample Period
Panel A: Inflation													
Canada	582	0.33	0.37	-0.86	2.31	0.67	1.90	132.48***	94.70***	418.46***	3.76*	23.08***	Feb 1968-Aug 2016
France	583	0.35	0.35	-0.45	1.80	0.88	0.41	80.08***	403.48***	1464.58***	234.04***	269.68***	Feb 1968-Aug 2016
Germany	583	0.21	0.25	-0.71	1.54	0.68	2.23	167.32***	67.04***	241.92***	9.84***	15.15***	Feb 1968-Aug 2016
Italy	583	0.51	0.51	-0.39	2.72	1.61	2.92	462.50***	412.13***	1496.03***	195.87***	218.32***	Feb 1968-Aug 2016
Japan	582	0.22	0.49	-0.65	3.61	2.46	10.48	3276.18***	136.11***	452.68***	191.38***	196.14***	Feb 1968-Aug 2016
UK	583	0.44	0.51	-0.63	4.48	2.56	12.00	4171.21***	278.93***	960.55***	44.68***	60.06***	Feb 1968-Aug 2016
US	583	0.33	0.33	-1.79	1.79	0.09	4.36	468.80***	235.01***	565.50***	129.19***	133.33***	Feb 1968-Aug 2016
Panel B: Gold price changes													
Spot	440	0.26	5.47	-26.12	36.86	0.76	6.79	897.92***	0.06	10.23**	8.57***	25.27***	Dec 1979-Aug 2016
1-month	440	0.26	5.41	-26.13	29.98	0.49	4.61	413.44***	0.00	11.71**	19.96***	21.57***	Dec 1979-Aug 2016
2-month	440	0.25	5.42	-26.18	30.08	0.50	4.71	431.82***	0.00	11.44**	18.48***	19.87***	Dec 1979-Aug 2016
3-month	440	0.25	5.45	-24.70	29.06	0.49	4.10	330.28***	0.01	11.75**	26.91***	18.38***	Dec 1979-Aug 2016
4-month	440	0.25	5.48	-25.14	29.57	0.48	4.27	357.16***	0.00	10.50**	24.23***	19.55***	Dec 1979-Aug 2016
5-month	440	0.24	5.47	-25.56	30.03	0.47	4.36	370.75***	0.00	11.99**	25.97***	17.97***	Dec 1979-Aug 2016
6-month	440	0.24	5.48	-25.94	30.42	0.47	4.48	389.50***	0.01	12.23**	25.73***	18.19***	Dec 1979-Aug 2016
7-month	440	0.23	5.50	-26.31	30.79	0.46	4.59	407.48***	0.02	12.43**	25.32***	18.65***	Dec 1979-Aug 2016
8-month	440	0.23	5.51	-26.65	30.95	0.46	4.66	419.41***	0.04	12.73**	25.43***	19.09***	Dec 1979-Aug 2016
9-month	440	0.23	5.52	-26.98	31.11	0.45	4.72	428.45***	0.07	13.04**	25.96***	19.79***	Dec 1979-Aug 2016
10-month	440	0.22	5.53	-27.23	31.27	0.45	4.79	441.77***	0.11	13.28**	26.43***	20.11***	Dec 1979-Aug 2016
11-month	440	0.22	5.55	-27.46	31.44	0.45	4.86	454.12***	0.17	14.07**	27.04***	21.02***	Dec 1979-Aug 2016
12-month	440	0.26	5.47	-26.12	36.86	0.76	6.79	897.92***	0.06	10.23**	8.57***	25.27***	Dec 1979-Aug 2016

Note: The table reports the descriptive statistics for seasonally adjusted monthly inflation rates (%) for the G7 countries and spot and futures (1- to 12-month) gold market price changes (%). Sample periods start at the period given in the last column of the table and end on August 2016 at a monthly frequency with *n* observations for each series. In addition to the mean, standard deviation (S.D.), minimum (min), maximum (max), skewness, and kurtosis statistics, the table reports the Jarque-Bera normality test (JB), Ljung-Box first [O(1)] and fourth [O(4)] autocorrelation tests, and first- [ARCH(1)] and fourth-order [ARCH(4)] Lagrange multiplier (LM) tests for the autoregressive conditional heteroskedasticity (ARCH). The asterisks *** ** and * represent significance at the 1%, 5%, and 10% levels, respectively.



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Although our aim is to investigate the causality in quantiles in both the mean and variance from inflation to gold market price changes, for the sake of completeness, we perform standard linear Granger causality tests using a linear VAR model. The results of the linear Granger causality tests are reported in columns 2 through 8 of Table 2. The resulting F -statistic given in columns (2) and (7) of Table 2 for the null hypothesis that inflation does not Granger-cause gold market price changes is less than 1.4 for Canada and 1.7 for the UK, respectively. The findings from the linear causality test indicate that there is no evidence of predictability from inflation to gold market price changes, even at a five percent significance level, for Canada and the UK. In column (3) of Table 2, we see that the null hypothesis that inflation does not Granger-cause gold spot gold market price changes is rejected for France only. However, the null hypothesis that inflation does not Granger-cause gold futures markets at various maturities (1-month, 2-month, and 3-month) is rejected at a 5% significance level and nine additional maturities (4-month, 5-month, 6-month, 9-month, 10-month, 11-month, and 12-month) at a 10% significance level. Linear Granger causality results for France indicate that inflation has predictive power for gold futures market, with the exception of the spot market. Similar to the findings for France, the evidence from column (4) in Table 2 for Germany indicate that the null hypothesis that inflation does not have predictive power for gold futures market at all maturities is rejected at the 10% significance level, while the null hypothesis that inflation does not have predictive power for gold spot market is not rejected at the 5% significance level. Contrary to the results obtained for France and Germany, the findings reported in Table 2 for Italy, Japan and the US show that the null hypothesis that inflation does not have predictive power for gold spot market is not rejected at the 5% significance level for these countries, while the null hypothesis that inflation does not have predictive power for gold spot market is rejected at the 1% significance level for Italy and Japan and at the 10% significance level for the US. The evidence shown in Table 2 can be summarized as follows. First, there is uniform and strong evidence at the 5% significance level for the power of inflation to predict gold market price changes for Canada, Italy, Japan, the UK and the US. Second, there is weak evidence of predictability from inflation to gold for Germany, while the results provide strong evidence of predictability from inflation to gold for France, with the exception of the spot market for both countries.

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Table 2. Linear Granger causality tests of the null hypothesis that inflation does not Granger-cause gold market price changes for each G7 country

Gold market	Canada	France	Germany	Italy	Japan	UK	US
Spot	1.3635	0.6068	0.0418	10.2515 ***	16.3884 ***	1.1419	3.6561 *
1-month	0.6440	5.2767 **	5.2326 **	0.7534	0.0290	1.6873	0.26551
2-month	0.6324	3.9849 **	5.3823 **	0.2689	0.0885	0.7538	0.07771
3-month	0.6146	4.0236 **	5.3173 **	0.2624	0.0890	0.6821	0.07201
4-month	0.6137	3.7694 *	5.4752 **	0.1435	0.1107	0.7929	0.03531
5-month	0.5985	3.7883 *	5.9056 **	0.1745	0.0599	1.0480	0.03601
6-month	0.6472	3.6839 *	5.3376 **	0.1756	0.0788	0.8026	0.03221
7-month	0.6663	3.6202 *	5.2710 **	0.1860	0.0634	0.8163	0.02791
8-month	0.6766	3.5313 *	5.1322 **	0.1959	0.0496	0.8330	0.02491
9-month	0.6932	3.4200 *	5.0837 **	0.1932	0.0356	0.8407	0.02131
10-month	0.6979	3.3027 *	4.9562 **	0.1916	0.0259	0.8380	0.01581
11-month	0.7072	3.1956 *	4.8813 **	0.1857	0.0183	0.8533	0.01241
12-month	0.7187	2.9975 *	4.8197 **	0.1808	0.0116	0.8592	0.01301

Note: The table reports the F-statistic for the no-Granger-causality restrictions imposed on a linear vector autoregressive (VAR) model under the null hypothesis H_0 . The order of the VAR is selected by the Bayesian information criterion (BIC). In the table, ***, **, and * indicate rejection of the null hypothesis of no Granger causality at the 1%, 5%, and 10% significance levels, respectively.

To justify the use of the nonparametric quantile-in-causality approach, we statistically examine the nonlinearity dynamics in the relationship between the inflation and gold price change series. We apply the Brock et al. (1996, BDS) test on the residuals of the gold price equation in a VAR(1) model of inflation and gold market price changes. The results given in Table 3 indicate strong evidence at the highest level of significance for the rejection of the null hypothesis of i.i.d. residuals at various embedding dimensions (m). These findings provide strong evidence of nonlinearity in gold market price changes. Thus, given the nonlinearity detected by the BDS test, the Granger causality tests based on a linear framework are likely to suffer from misspecification. Therefore, the results of the linear Granger causality test cannot be deemed reliable. Given the strong evidence of nonlinearity, we turn our attention to the nonparametric causality-in-quantiles test, which is robust to functional misspecification, as it is a nonparametric (i.e., data-driven) approach.

In Figures 3-9, we present the findings of the nonparametric causality-in-quantiles tests for causality from the inflation series to the gold spot and futures price change series for Canada, France, Germany, Italy, Japan, the UK and the US, respectively. The figures report the results for both the causality in mean and causality in variance.

Table 3. [Brock et al. (1996)] BDS nonlinearity test

Gold Market	<i>m</i> =2	<i>m</i> =3	<i>m</i> =4	<i>m</i> =5	<i>m</i> =6
Canada					
Spot	8.0836***	10.5065***	12.6977***	15.6595***	18.8721***
1-month	6.1975***	8.2798***	10.9559***	13.5394***	14.9734***
2-month	5.8825***	7.7882***	10.2361***	12.4313***	13.9021***
3-month	5.7787***	7.4884***	9.6189***	11.4365***	12.9964***
4-month	5.8114***	7.7030***	9.3201***	11.1074***	13.2374***
5-month	4.9929***	6.8735***	8.6422***	10.8534***	13.4207***
6-month	5.2925***	7.0103***	8.8667***	11.0644***	13.6667***
7-month	4.9708***	6.7356***	8.6835***	10.6956***	12.9433***
8-month	4.9924***	6.7382***	8.7529***	11.0350***	13.5324***
9-month	4.6728***	6.5385***	8.7002***	10.7985***	12.9979***
10-month	4.6048***	6.4444***	8.7549***	11.0304***	13.2196***
11-month	4.4476***	6.1970***	8.1991***	10.5433***	12.2424***
12-month	4.0707***	6.0809***	8.1123***	10.4975***	12.3199***
France					
Spot	8.3298***	10.9926***	13.4937***	16.6212***	20.4718***
1-month	5.4320***	7.6419***	9.9553***	12.4475***	13.9305***
2-month	5.4702***	7.5190***	9.5300***	11.7862***	13.4536***
3-month	5.2162***	7.2850***	9.3257***	11.1762***	12.4493***
4-month	5.6760***	7.4833***	9.1383***	11.0351***	12.3833***
5-month	4.9013***	6.8987***	8.9659***	10.8379***	12.6651***
6-month	5.1902***	7.2348***	9.1539***	10.8733***	12.5182***
7-month	5.0249***	7.0332***	9.0641***	10.9055***	12.1858***
8-month	4.7848***	6.9448***	8.9391***	11.0990***	12.4099***
9-month	4.5668***	6.6630***	8.9368***	11.2932***	12.9471***
10-month	4.3260***	6.2412***	8.3188***	10.3793***	12.1480***
11-month	4.0323***	6.0421***	8.0468***	10.2539***	12.2687***
Germany					
Spot	7.9761***	10.5108***	12.8290***	15.6873***	19.1868***
1-month	6.8318***	9.1409***	11.2467***	13.4961***	15.1028***
2-month	6.4651***	8.6119***	10.5781***	12.5115***	14.2459***
3-month	6.3598***	8.1621***	10.0171***	12.2352***	14.7125***
4-month	6.4040***	8.3843***	10.1910***	11.6455***	13.4460***
5-month	5.7087***	7.7907***	9.7054***	11.5255***	13.6768***
6-month	5.7898***	8.0462***	9.8211***	11.3624***	13.3372***
7-month	5.6663***	7.9580***	9.8827***	11.4297***	13.0755***
8-month	5.4046***	7.6244***	9.6103***	11.3137***	13.1897***
9-month	5.1047***	7.1263***	9.2054***	10.7688***	13.3050***
10-month	4.7911***	6.8609***	8.7956***	10.2797***	12.4558***
11-month	4.5640***	6.6769***	8.5657***	10.0269***	11.7121***
12-month	4.3772***	6.5137***	8.3875***	9.9233***	11.4546***
Italy					
Spot	7.7253***	9.8736***	12.3246***	15.7227***	19.7286***
1-month	6.4129***	8.5105***	10.7158***	13.2622***	15.1096***
2-month	6.1743***	7.9048***	10.1619***	12.3284***	13.9940***
3-month	5.9442***	7.5319***	9.6356***	11.4473***	12.4579***
4-month	5.8652***	7.5163***	9.1236***	11.0007***	13.3274***
5-month	5.0732***	7.0320***	8.9645***	11.1096***	13.0371***
6-month	5.3783***	7.2362***	9.0764***	11.2165***	12.8696***
7-month	5.2617***	6.9857***	9.0576***	11.0024***	12.8730***
8-month	5.1356***	6.9372***	8.9649***	11.0174***	13.2294***
9-month	4.9222***	6.6279***	8.7792***	10.7791***	13.1932***
10-month	4.8470***	6.4612***	8.5170***	10.5085***	12.4302***
11-month	4.6619***	6.3764***	8.4233***	10.2991***	12.1910***
12-month	4.2294***	6.0705***	8.0221***	9.8652***	11.2196***

Note: The entries indicate the BDS test based on the residuals of the gold equation in the VAR for various inflation and gold return VAR models. *m* denotes the embedding dimension of the BDS test. In this table, ***, ** and * indicate rejection of the null hypothesis that residuals are i.i.d. at the 1%, 5%, and 10% significance levels, respectively.

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Table 3. (continued)

Gold Market	<i>m</i> =2	<i>m</i> =3	<i>m</i> =4	<i>m</i> =5	<i>m</i> =6
Japan					
Spot	6.8498***	8.8987***	10.9353***	13.9018***	16.8448***
1-month	6.4287***	8.4376***	10.7389***	12.9873***	14.6144***
2-month	5.9545***	8.0622***	10.3745***	12.5752***	15.6516***
3-month	5.6100***	7.4597***	9.7847***	12.0670***	14.2110***
4-month	5.8312***	7.7800***	9.4764***	11.4818***	13.2727***
5-month	4.8679***	6.8164***	8.6668***	10.7947***	12.8730***
6-month	5.2969***	7.2322***	9.0392***	11.0907***	13.1450***
7-month	5.0469***	6.9671***	9.0317***	11.1516***	13.3873***
8-month	5.0610***	7.0447***	9.1742***	11.0815***	12.9707***
9-month	4.7300***	6.6428***	8.9637***	11.1402***	13.4120***
10-month	4.7455***	6.4441***	8.6046***	10.7929***	13.2932***
11-month	4.4347***	6.1937***	8.3213***	10.5489***	12.5651***
12-month	4.2400***	6.1070***	8.0782***	10.2009***	11.7597***
UK					
Spot	7.7336***	10.4951***	12.8655***	16.0190***	19.7654***
1-month	6.5859***	8.3074***	10.2012***	12.0538***	13.7528***
2-month	6.1401***	7.6873***	9.5142***	11.4746***	13.5932***
3-month	5.9953***	7.5762***	9.4583***	11.3192***	13.4744***
4-month	6.0534***	7.6878***	9.0074***	10.6014***	11.9660***
5-month	5.3932***	7.1299***	8.7506***	10.2304***	11.0223***
6-month	5.4378***	7.0442***	8.4717***	9.9852***	11.1373***
7-month	5.3323***	7.0190***	8.7308***	10.1802***	11.8988***
8-month	5.1973***	6.9048***	8.7179***	10.2776***	11.3407***
9-month	4.9637***	6.6564***	8.4995***	10.0763***	11.1176***
10-month	4.7487***	6.4633***	8.2150***	9.8007***	10.5389***
11-month	4.5948***	6.3607***	8.2939***	10.3098***	11.6561***
12-month	4.3849***	6.2629***	8.1836***	10.2957***	12.4576***
US					
Spot	8.3255***	10.6854***	13.3487***	17.0143***	20.7251***
1-month	6.4092***	8.5255***	10.9538***	13.7189***	15.9628***
2-month	6.2163***	8.1281***	10.3805***	12.8860***	15.6961***
3-month	5.8329***	7.4664***	9.6421***	11.5647***	13.3464***
4-month	5.8727***	7.8253***	9.3630***	11.6338***	14.3397***
5-month	5.0365***	7.0218***	8.8433***	11.0393***	13.2542***
6-month	5.2591***	7.0063***	8.8782***	11.1851***	13.0522***
7-month	5.2144***	7.0874***	9.0959***	11.1356***	12.9572***
8-month	5.1351***	6.9566***	9.1118***	11.4776***	13.7684***
9-month	4.8989***	6.6872***	8.9497***	11.0861***	13.5248***
10-month	4.7997***	6.5057***	8.6778***	10.8184***	13.0355***
11-month	4.6581***	6.3413***	8.5116***	10.7238***	12.9202***
12-month	8.3255***	10.6854***	13.3487***	17.0143***	20.7251***

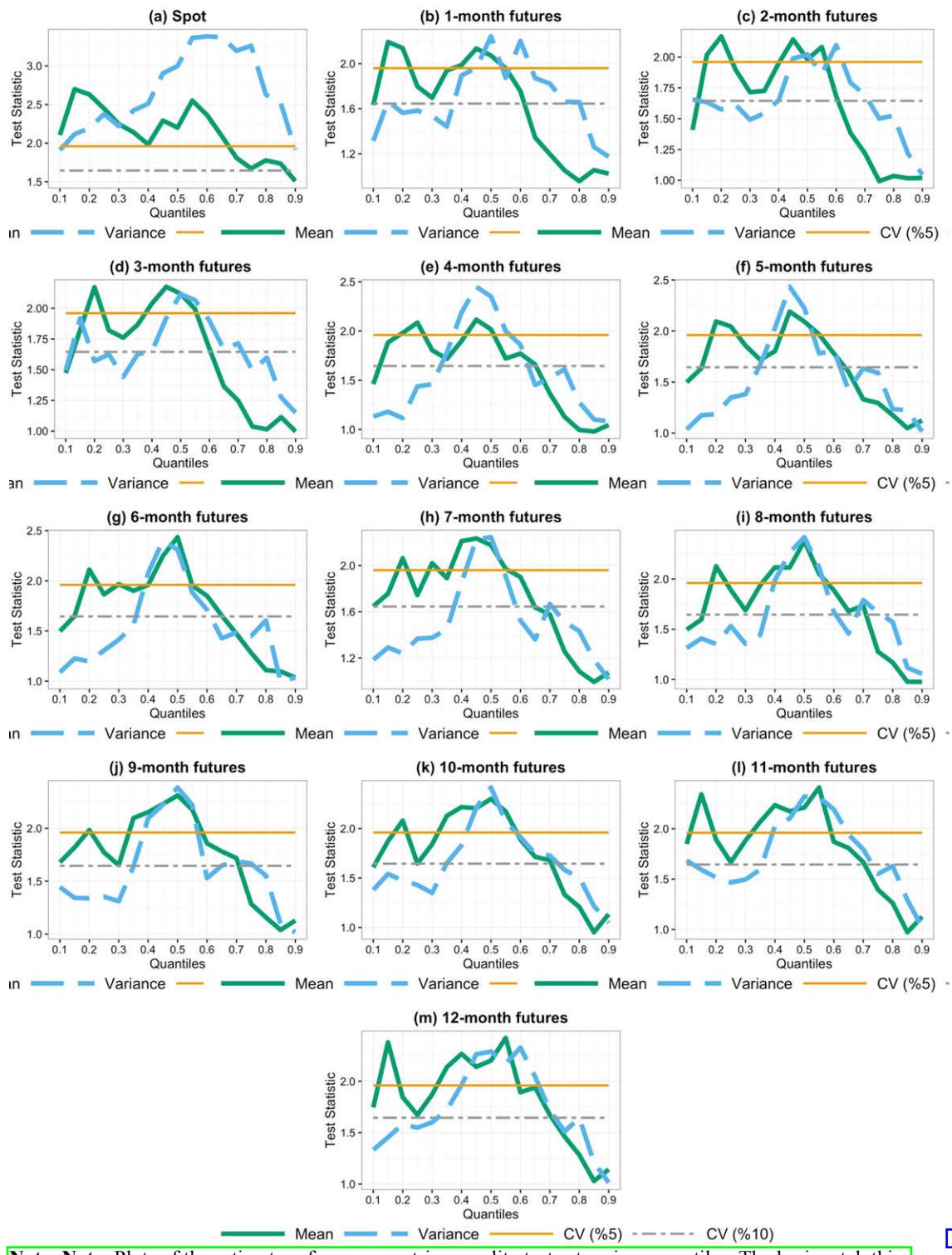
□

In each figure, the vertical axis includes the test statistic corresponding to the null hypothesis that inflation does not Granger-cause gold market price changes, and the horizontal axis includes the quantiles. The 5% critical value is 1.96, and the 10% critical value is 1.64. In these figures, the horizontal solid lines and the horizontal two-dashed lines represent the 5% and 10% critical values, respectively.

□

□

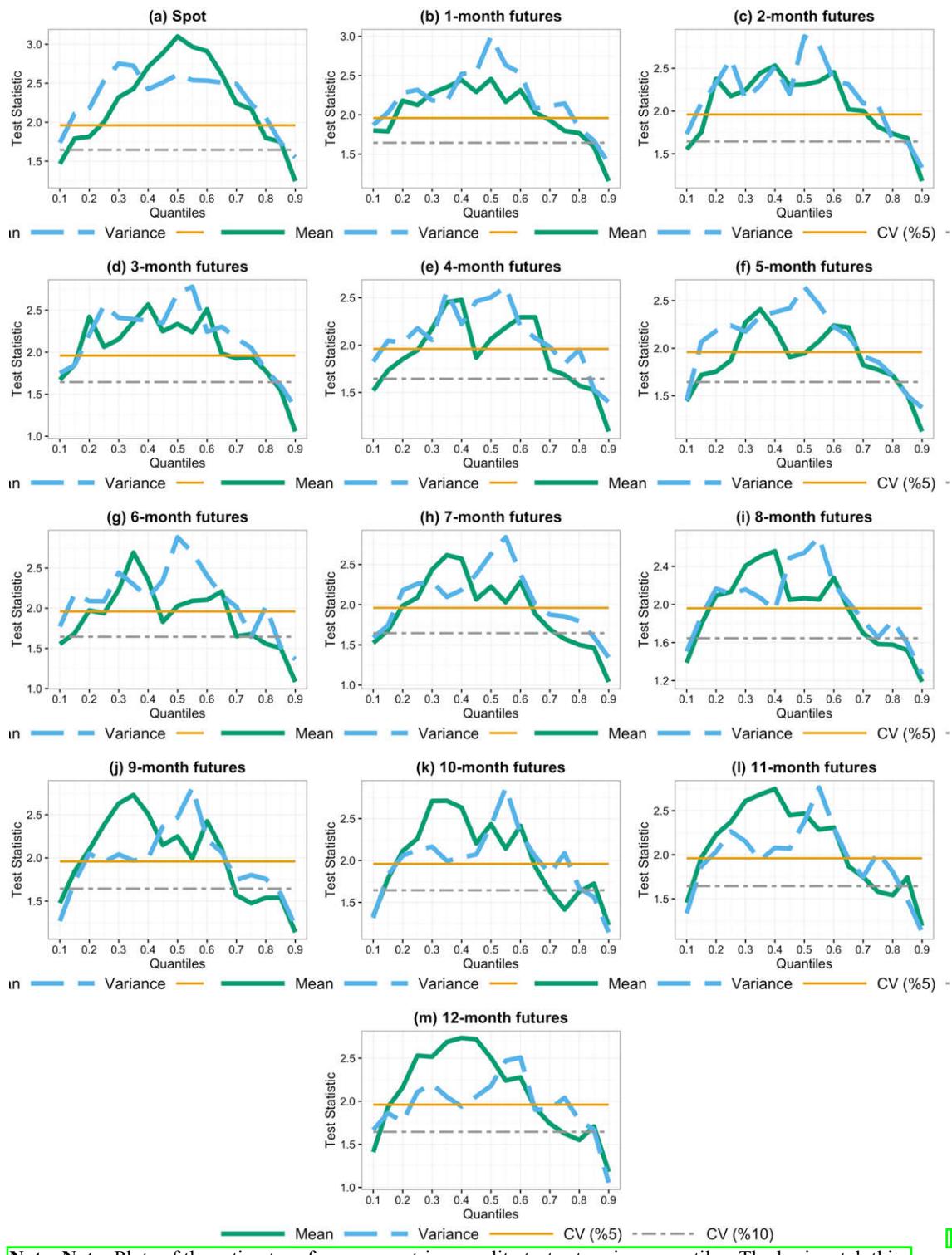
□



Note: Plots of the estimates of nonparametric causality tests at various quantiles. The horizontal, thin, solid lines and the thin, two-dashed lines represent the 5% and 10% critical values, respectively.

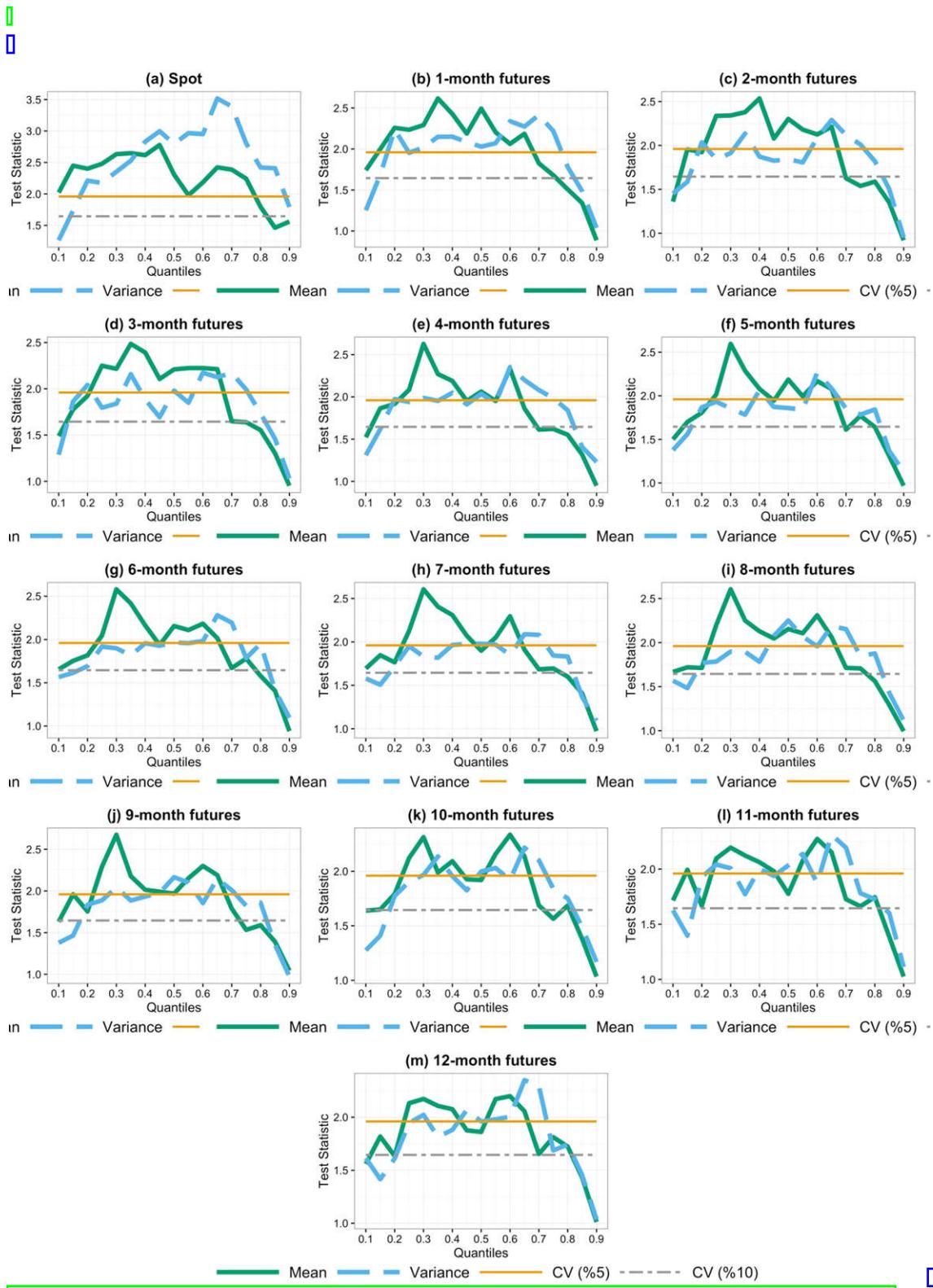
Figure 4. Causality in mean and variance from inflation to the gold market series for France





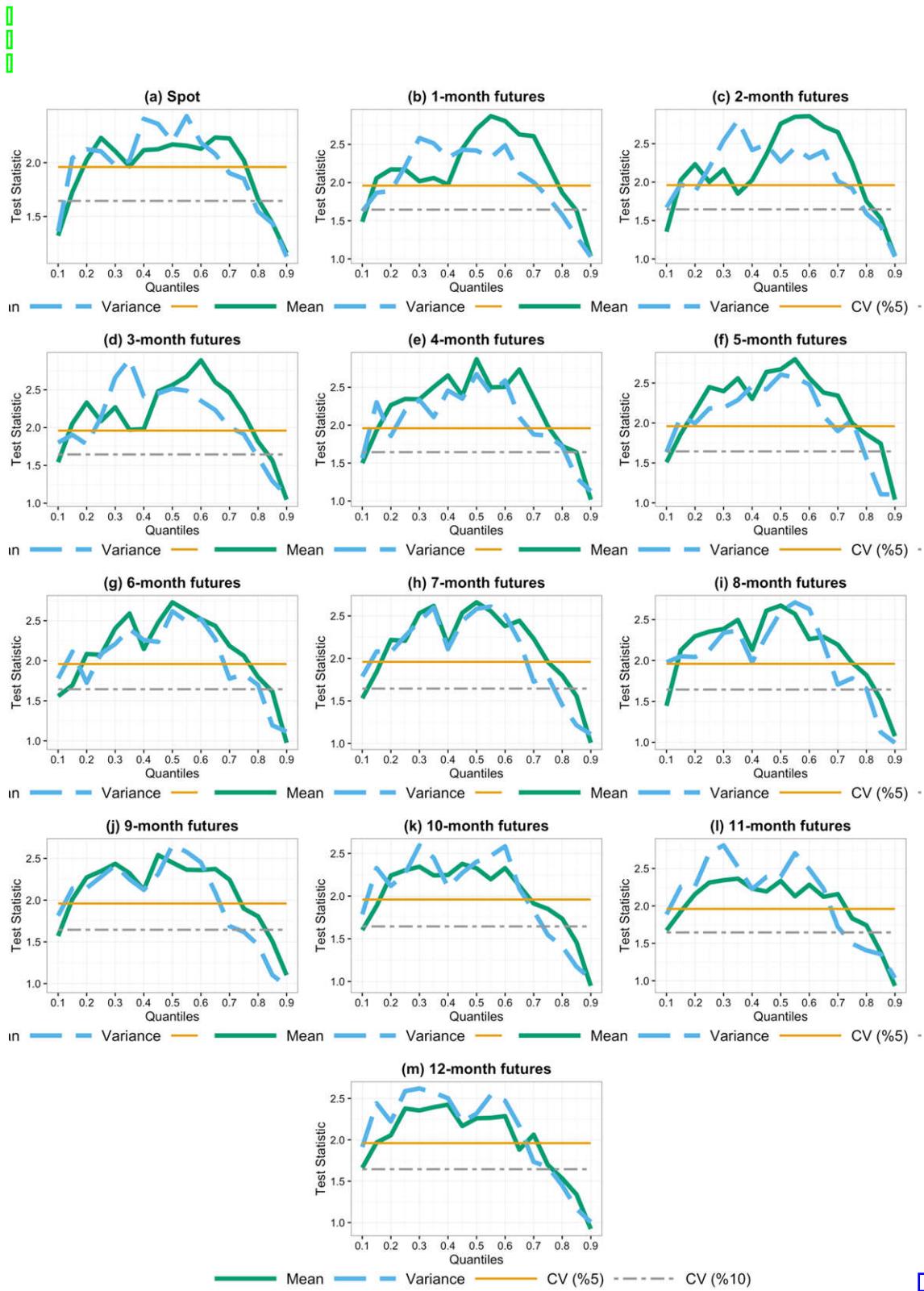
Note: Plots of the estimates of nonparametric causality tests at various quantiles. The horizontal, thin, solid lines and the thin, two-dashed lines represent the 5% and 10% critical values, respectively.

Figure 5. Causality in mean and variance from inflation to the gold market series for Germany



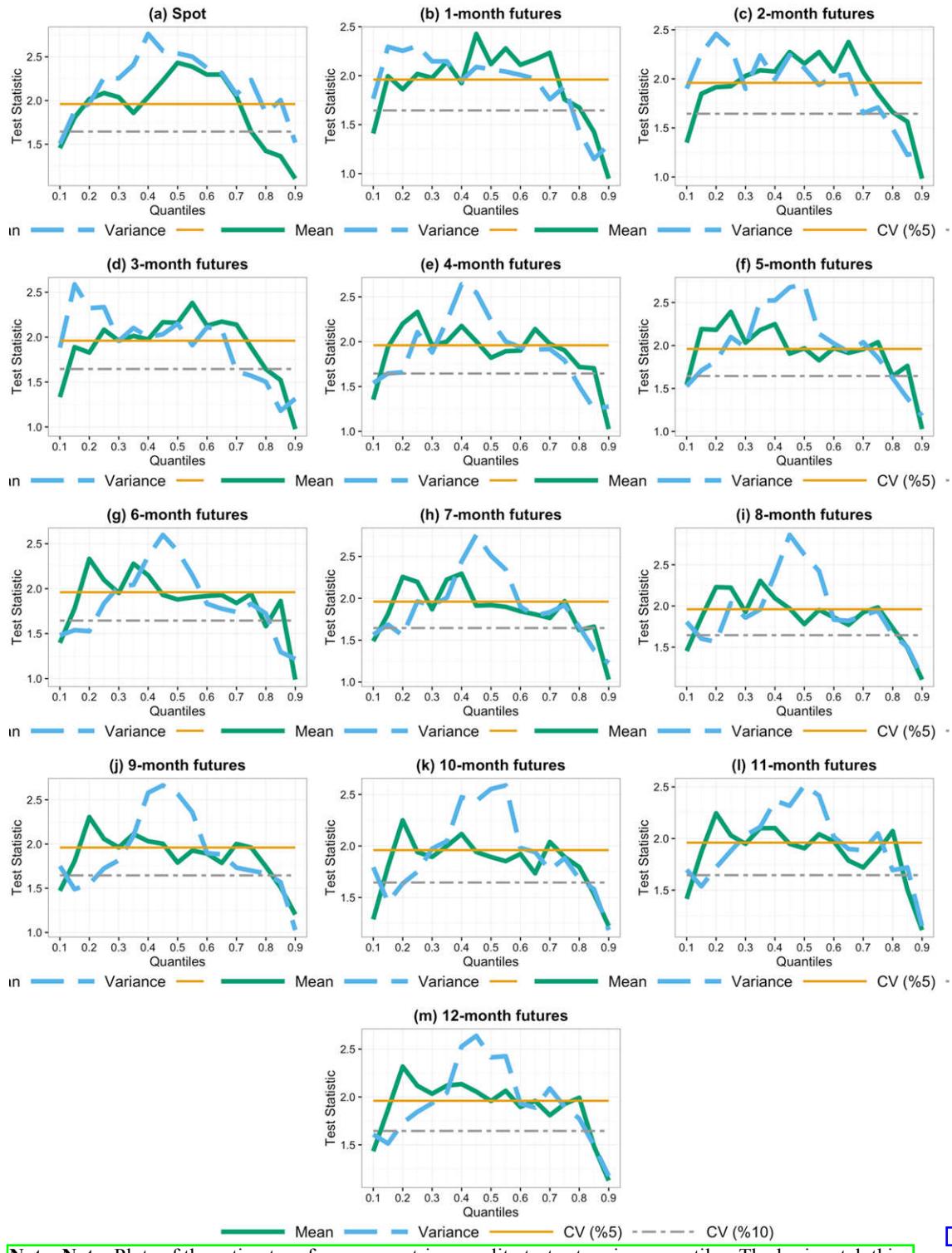
Note: Plots of the estimates of nonparametric causality tests at various quantiles. The horizontal, thin, solid lines and the thin, two-dashed lines represent the 5% and 10% critical values, respectively.

Figure 6. Causality in mean and variance from inflation to the gold market series for Italy



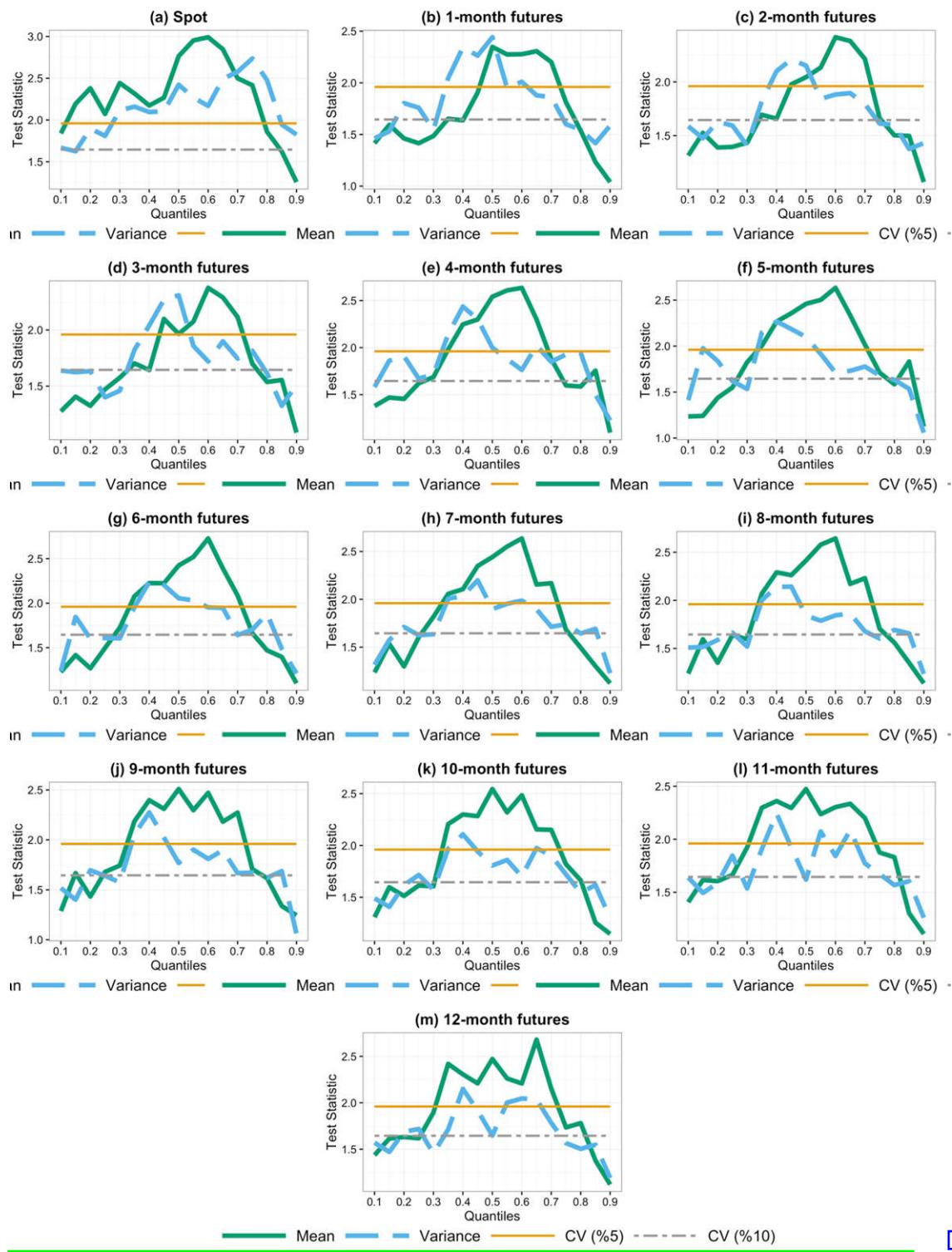
Note: Plots of the estimates of nonparametric causality tests at various quantiles. The horizontal, thin, solid lines and the thin, two-dashed lines represent the 5% and 10% critical values, respectively.

Figure 7. Causality in mean and variance from inflation to the gold market series for Japan



Note: Plots of the estimates of nonparametric causality tests at various quantiles. The horizontal, thin, solid lines and the thin, two-dashed lines represent the 5% and 10% critical values, respectively.

Figure 8. Causality in mean and variance from inflation to the gold market series for the UK



Note: Note: Plots of the estimates of nonparametric causality tests at various quantiles. The horizontal, thin solid lines and the thin, two-dashed lines represent the 5% and 10% critical values, respectively.

Figure 9. Causality in mean and variance from inflation to the gold market series for the US

I

The findings reported in Figures 3-9 indicate that the results relating to causality in variance have some clear differences from the results for the causality in mean. The evidence in Figure 3 indicates that there is causality in mean and variance from inflation to gold market price changes for Canada around the quantile ranges from 0.25 to 0.70. This result indicates that inflation has predictive power for gold market price changes in Canada. The results for Germany, Italy, Japan and the UK reported in Figures 5, 6, 7 and 8, respectively, are analogous to the findings for Canada. Thus, the evidence from Figures 5-8 shows that the null hypothesis that inflation has no predictive power for gold market price changes is rejected at the five percent significance level around the quantile range of 0.25 to 0.70, with the exceptions of the lower or upper quantiles for Germany, Italy, Japan, and the UK. This finding indicates that inflation has strong predictive power for gold market price changes.

I

For France, the results in Figure 4 indicate that the null hypothesis that inflation does not Granger-cause mean gold market price changes is rejected over the quantile range of 0.10 to 0.90, while the null hypothesis that inflation does not Granger-cause gold market price changes in variance is rejected over the quantile range of 0.10 to 0.65 for the gold spot market. These findings show that there is strong evidence of predictability from inflation to gold market price changes for the gold spot market. Contrary to the results for the gold spot market case for France, the findings from futures markets at all maturities show that there is weak evidence against the null hypothesis that inflation does not Granger-cause the mean and variance in gold market price changes. These estimates reveal weak evidence that inflation Granger-causes the mean and variance in gold market price changes for gold futures markets at all maturities for France. The results for the US reported in Figure 9 are analogous to the results for France.

I

The results of causality-in-quantiles tests indicate that inflation has predictive power for gold market price changes. This should be the case because inflation and gold market price changes are strongly interdependent, as expected. Causality-in-quantiles tests indicate causality from the 0.20 quantile to the 0.70 quantile range in most cases. This result leads us to conclude that the changes in low- and high-price change ranges in the gold market are not associated with inflation. They may be linked to other factors, such as financial shocks and exchange market shocks. Our findings support the widely regarded idea that gold is a hedge

against inflation, but they also show when the hedging role of gold breaks down. Since there is causality in variance as in mean (price changes), it is possible to argue that the impact of inflation on gold market volatility may lead to policy uncertainty and consequently increase the cost of policy implementation. It also increases uncertainty among investors and leads to higher risk premiums on gold positions. Our findings may help investors to ensure better asset allocation in inflationary environments. In terms of the continued hoarding of general public investment in gold and large gold reserves held by central banks, the findings of this study suggest that investors should be advised to monitor and be cautious, as their confidence might be eroded by structural changes. The evidence from this study highlights the need to account for nonlinearities when modelling the dynamics of the gold price-inflation relationship. Moreover, the analysis in this study can be extended to gold hedging strategies in inflationary environments.

□

4. Conclusion

There exists an extensive body of literature investigating the role of inflation in predicting developments in gold markets. We expand upon this literature by investigating the impact of inflation on gold market price changes and gold market volatility using a novel nonparametric causality-in-quantiles approach. In the paper, we use monthly data for the sample from December 1979 to August 2016. To assess consistency, we first apply the standard linear Granger causality test. Linear tests might be misleading because they ignore nonlinear dependence and therefore cannot detect causality from inflation to gold market price changes when nonlinearities exist. The result of the nonlinearity test of Brock et al. (1996) indicates that inflation and gold return series have a strong nonlinear dependence. The evidence of nonlinearity suggests that the linear framework to test causality is in fact misspecified, with unreliable results from these tests. Consequently, we use a novel nonparametric causality-in-quantiles test that combines elements of the test for nonlinear causality of the k -th order by Nishiyama et al. (2011) with the causality-in-quantiles test developed by Jeong et al. (2012). The causality-in-quantiles approach allows us to test for causality in both the mean and the variance (volatility). This approach adds additional power to the test when causality in the conditional mean does not exist yet higher-order interdependencies do. Causality-in-quantiles tests find causality from inflation to all spot and futures gold markets, as expected. Intuitively, this should be the case because spot and futures gold markets are highly correlated. Causality-in-quantiles tests usually find causality from the 0.20 quantile to the quantile 0.70, meaning

that very low- and high-gold price movements in gold markets are not related to inflation. These changes should be related other factors, such as financial shocks and exchange market shocks. We find support that gold serves as a hedge against inflation, a common assumption everywhere in the world. Our findings, however, show that the hedging power of gold does not hold at quantiles lower than 0.20 and higher than 0.70, which corresponds very low- and high-level gold market price change ranges. Causality is not only in the mean (price changes) but also in variance, meaning that inflation has an impact on gold market volatility. In general, our results, via the nexus between inflation and the gold market, highlight the importance of detecting and modelling nonlinearity when analysing predictability via causal relationships.



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