ISSN 1791-3144

University of Macedonia
Department of Economics

Discussion Paper Series

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Discussion Paper No. 3/2015

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http://www.uom.gr/index.php?newlang=eng&tmima=3&categorymenu=2

Are Gold and Silver a Hedge against Inflation? A Two Century Perspective

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Abstract

This study examines the long-run hedging ability of gold and silver prices against alternative measures of consumer price index for the UK and the US. We employ a dataset that spans from 1791 to 2010, and both a time-invariant and a time-varying cointegration framework. We find that gold can at least fully hedge headline, expected and core CPI in the long-run. This ability tends to be stronger when we allow for the long term dynamics to vary over time. The inflation hedging ability of gold is on average higher in the US compared to the UK. Silver does not hedge US consumer prices albeit evidence emerges in favor of a time-varying long-run relationship in the UK.

Keywords: gold prices, silver prices, inflation hedge, time-varying cointegration.

1. Introduction

Gold and silver have played a major role in the history of money and monetary policy. They have traditionally acted as medium of exchange, store of wealth, and a unit of value (Goodman, 1956; Solt and Swanson, 1981). In contrast to many other multi-faceted commodities, they are durable, relatively transportable, universally acceptable and easily authenticated. Gold as the most acclaimed precious metal in human history, still plays a pronounced role as a store of value especially in times of uncertainty. This feature stems from the 'flight to quality' behavior of investors who purchase gold in search for safer assets (Baur and Lucey, 2010). As precious metals represent claims to real rather than nominal assets, under the Fisher (1930)'s framework, gold and silver are expected to hedge against inflation.[1] An expected increase in consumer price level may induce individuals to convert their current liquid assets into gold, influencing

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[1]The proposition that *ex ante* nominal asset returns contain the market's perception of anticipated inflation rates can be applied to all assets in efficient markets, also known as Generalized Fisher Effect (GFE, henceforth) (see, Jaffe and Mandelker, 1976).

its price (Fortune, 1987).[2] Therefore, gold and silver prices could effectively gauge inflation expectations since, commodity prices are generally considered to be able to incorporate new information faster than consumer prices (Mahdavi and Zhou, 1997).

To the best of our knowledge, historical data have not being employed so far to examine the long-run (*LR*) hedging ability of gold and silver. This article fills this gap by looking at over 200 years of annually data for the UK and the US gold and silver markets. The renewed attention for the role of precious metals, as fundamental investment strategy against the eroding impact of inflation, further reinforces the goals of this study: (i) assess the historical role of gold and silver prices as a hedge against headline, expected and core CPI measures and (ii) examine the hedging ability of these two precious metals in a time-invariant (TI) and a time-varying (TV) cointegration framework that allows for nonlinear adjustment and the smooth evolution of the long-run relationship.

Laurent (1994) notes that during the 1800 to 1992 period, the price of gold and the general level of prices (wholesale prices) in the US have corresponded quite closely. Jastram (1978) indicates that the study of gold in the US is a logical companion piece to the study of UK, given that economic institutions are akin and common factors influence their commerce and finance. London was the undisputed center of the world capital markets during the gold standard, since the Bank of England could exert a powerful influence on the money supplies and price levels of other gold-standard countries (Bordo and Schwartz, 1994). Moreover, the US has been a prime mover in silver markets since the last quarter of the nineteenth century (Jastram, 1981). Concerning inflation, Siegel (2008) ascertains a similar pattern between the US and the UK consumer price level in the last 150 years, which is characterized by a significant overall inflation until World War II and protracted inflation later. Thus, the UK and the US can be considered as the most preferable cases for examining the historical long-run hedging ability of gold and silver.

Our paper is broadly related to a literature that used cointegration to study the hedging abilities of an asset in the long-run (Ely and Robinson, 1997; Anari and Kolari, 2001). Research into the gold/CPI relationship, though, produced mixed evidence. Garner (1995), Mahdani and Zhou (1997) and Tully and Lucey (2007) have documented an insignificant *LR* relationship between gold price and CPI, while Gosh et al. (2004) and Worthington and Pahlavani (2007) assert a significant positive relationship in the US.[3] On the other hand, the empirical literature for the hedging ability of silver is

[2]The theoretical argument developed in Fortune (1987) emphasized the substitution effect rather than the wealth effect. The author developed an equilibrium model, focusing on the US demand side of gold.

[3]See also Chua and Woodward (1982), Kaufmann and Winters (1989) and Tkacz (2007) who document

less extensive. Adrangi et al. (2003) argue that investment in gold and silver may be reliable inflation hedges in both the short- and long-term. McCown and Zimmerman (2006, 2007) provide evidence in favor of the hedging ability of gold and silver against inflation risk, especially over longer time horizons. Aggarwal and Lucey (2007) study psychological barriers in gold prices. A recent study of Wang et al. (2011) observes that time and market selection are the keys to inflation hedge. They employ a threshold cointegration framework and find that the low cross elasticity, the incomplete price adjustment and the short-run rigidity of the price adjustment between gold price and CPI might eliminate the inflation hedge ability of gold.

The 2007-09 financial crisis, the rise in volatility of commodity prices in conjunction with the tendency of central banks to become net buyers of bullion have revived the discussion on the role of precious metals.[4] In March 2011, Chatham House Gold Taskforce was founded in order to captivate the multiple role of gold (including that of hedge against inflation) as a means of enhancing the performance of the international monetary system.[5] One widely held argument for a renewed role for gold is that its countercyclical qualities can serve as a hedge against specific risks, such as bouts of inflation or financial contagion. However few would argue that a return of gold as an anchor in the international monetary system is feasible.[6] Nevertheless, from an investor's point of view, an examination of whether gold and silver prices historically maintain their value relative to consumer prices becomes increasingly important for several reasons.

From a practical perspective, many investors hold precious metals over long holding periods. Therefore, it is crucial to examine whether gold and silver prices move together with consumers prices over longer horizons. This applies to both long-term institutional and individual investors, for whom real-term capital preservation is a primary objective. In addition, the puzzling results of previous studies as well as the provided evidence for instabilities in precious metals and goods prices further rein-

a positive relationship between gold and inflation in the US.

[4]Speaking in the Financial Times before the G20's core reform agenda in Cannes, Robert Zoellick, the president of the World Bank, has stated that a new monetary system should 'consider employing gold as an international reference point of market expectations about inflation, deflation and future currency values'(see, Zoellick, R. 'The G20 must look beyond Bretton Woods II', Financial Times, 7 November 2010).

[5]Between 1929 and 1931, Chatham House convened a special Study Group having John Maynard Keynes as a member, in order to examine the problems arising from the post-war international monetary settlement, which contributed to the Great Depression and ultimately led to the suspension of the Gold Standard by the British government in September 1931. Available at: http://www.chathamhouse.org/publications/papers/view/178235

[6]The term anchor refers to whether gold has a role in being tied to or linked with the expansion or contraction of the global monetary base.

force the case for employing a time-varying approach (see, e.g. Beckmann and Czudaj, 2013; Batten et al., 2014). In our analysis, the *LR* coefficients quantify the intensity of the relationship between the two precious metals and alternative consumer price measures. Gold and silver could be a poor hedge against inflation in the short-term, but as the investment horizon increases they may provide adequate *LR* hedging properties. Furthermore, while investors and central banks have been buying gold in order to protect themselves against inflation risk, less attention has been given to silver. Silver, as one of the most attractive naturally occurring elements, may also provide inflation hedging properties.[7] Lastly, precious metals and consumers prices are both known to be integrated processes, thus estimating regressions in terms of their first (or higher order) differences implies partial loss of valuable long-run information (Anari and Kolari, 2001).

With these concerns in mind, this study examines the generalized Fisher effect using over two centuries of data for gold, silver and consumer prices. We employ time-invariant and time-varying cointegration analysis, that allows us to utilize the long-run information and account for different regimes. The key findings of the paper are as follows: (i) the real price of gold and silver is stationary when we account for structural breaks, (ii) we get moderate (strong) evidence of time-invariant (time-varying) cointegration between the precious metal prices and alternative CPI's for the US and the UK, (iii) the average *LR* betas for gold are above unity indicating superior hedging ability,0 (iv) the average long-run beta for gold is higher in the US compared to the UK, (v) more stability is observed in the long-run hedging ability of gold vs expected inflation during the last decades and (vi) the long-run relationship between silver and CPI in the UK emerges only when we consider the case of time-varying cointegration.

The remaining parts are as follows: Section 2 describes the data, Section 3 presents the methodology. The results are discussed in Section 4 and Section 5 concludes.

2. Data

The empirical analysis is conducted using annual data on consumer prices, gold prices and silver prices in the UK and US over the period 1791 to 2010.[8] The sample period for silver prices starts in 1792 for both countries. The inflation series are obtained from Reinhart and Rogoff (2011) (RR series).[9] Gold and silver prices are

[7]The time period that central banks viewed silver on par with gold as a reserve can be traced back to 1923. The St Louis Fed's report on this issue is available at: http://fraser.stlouisfed.org/docs/publications/FRB/pages/1920-1924/43097_1920-1924.pdf

[8]A detailed description of the data is given in the Appendix.

[9]Note that the base of the CPI was set to 100 in 1791.

obtained from Officer and Williamson (2011) and the Kitco Metals Inc, respectively.[10] Following Bekaert and Wang (2010), the precious metals prices were converted into local currency (USD $ and GBP £ per ounce), therefore, their hedging ability may also be due to currency movements, rather than to changes in their prices *per se*. The data (in logarithms) for the US and the UK are presented in Figures 1 and 2, respectively. An initial observation suggests that gold, silver and consumers price data went through many shifts over time, in both countries. We observe that for long periods in the 19th and 20th century gold and silver prices remain constant, reflecting the nominal price rigidity under periods of *monometallic* or *bimetallic* regimes.

In order to extract expected CPI measures, we employ two methodologies: the linear Hodrick and Prescott (1980, henceforth HP) filter and the asymmetric band-pass filter proposed by Christiano and Fitzgerald (2003, henceforth CF).[11] Each of these filters produce a long-term trend component of a series that may then be used to examine the long-run relationship of the historical prices of gold and silver and the expected consumer price level.[12]

We employ two different estimates for core CPI: the exponentially smoothed core inflation estimator proposed by Cogley (2002) and a wavelet method proposed more recently by Dowd et al. (2011).[13] The later compare the wavelet-based core measures against a number of alternative measures (including Congley's low pass filter) and conclude that the former generally performs better. We utilize both the single exponential smoothing (*ses*) and two core inflation measures based on wavelets analysis namely *d*4, *la*8.[14]

Figure 3 summarizes the different measures. For the CPI series (headline, expected and core) and the nominal gold and silver prices of both countries under investigation,

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[10]http://www.measuringworth.com/gold/ and http://www.kitco.com

[11]For the HP method, we follow Ravn and Uhlig (2002) by setting the penalty parameter equal to 6.25, for annual data.

[12]Ash et al. (2002), using long annual data, evaluate the usefulness of HP-filtered time series as a proxy for rational expectations applying a battery of tests for rationality. From their analysis of the US inflation data, they conclude that over the long period of 120 years the HP filter is weakly rational, being unbiased but inefficient. Orr et al. (1995) and Martin and Scarpetta (1999) also proxy inflation expectations with the HP filter. Smant (1998) argues that the HP filter generates a price series which is consistent with price expectations formed rationally indicating that HP filter incorporates a substantial element of 'perfect foresight' price expectations.

[13]Wavelet analysis involves the application of successive approximations to remove more and more high-frequency detail, or noise, and so help reveal an underlying signal (or shape) of the data. The procedure for the selection of plausible wavelet-based measures of core inflation is described in Dowd et al. (2011, pp. 521-522).

[14]The wavelets were calculated using the R package: "wavelets", covering four classes of wavelet transform filters: Daubechies, Least Asymetric, Best Localized and Coiflet. The prefixes for filters of these classes are *d*, *la*, *bl* and *c*, respectively. The optimal number of wavelets level was selected according the lowest value of the Shannon entropy (R package: "wavethresh").

we have conducted three unit root tests. The augmented Dickey and Fuller (1979, ADF henceforth), the Phillips and Perron (1988, PP henceforth), and the Ng and Perron's (2001, NP henceforth) all seem to suggest that the maximum order of integration (d) is 1. The results are reported in Table 2.

3. Methodology

In order to assess the underlying long-run dynamics, we employ two cointegration methodologies; the trace test proposed by Johansen (1995) and the Bierens and Martins (2010) time-varying (TV) vector error correction model, in which the cointegrating relationship varies smoothly over time and the adjustment can be nonlinear. Johansen (1995) cointegration procedure is restrictive in the sense that it assumes that the cointegrating vector is constant and the adjustment is linear.

The time-invariant Vector Error Correction model (TI-VECM) of order p, used to construct the Johansen tests can be written as:

$$\Delta Z_t = \Pi Z_{t-1} - \gamma_0 - \gamma_1 t + \sum_{j=1}^{p-1} \Gamma_j \Delta Z_{t-j} + \varepsilon_t, \quad \varepsilon_t \sim N_k(0, \Omega), \quad t=1,\ldots,T, \tag{1}$$

where Z_t is a $k \times 1$ vector of variables observed at time t, Ω and Γ_j, $j=1,\ldots,p-1$, are $k \times k$ matrices. Johansen (1995) used full information maximum likelihood cointegration analysis to test for a long-run relationship.[15] If cointegration exists, one can decompose $\Pi=\alpha\beta'$ and (1) can be rewritten as:

$$\Delta Z_t = \alpha(\beta' Z_{t-1} - \beta_0 - \beta_1 t) - \gamma_0 - \gamma_1 t + \sum_{j=1}^{p-1} \Gamma_j \Delta Z_{t-j} + \varepsilon_t, \tag{2}$$

where α is a $k \times r$ matrix of coefficients (where r is the cointegrating rank of the system), β is a $k \times r$ matrix of coefficients which defines the r cointegrating vectors in the system, β_0 is an $r \times 1$ vector of intercepts for the cointegrating vectors, β_1 is a $r \times 1$ vector of coefficients which allows for linear deterministic trends in the cointegrating vectors, γ_0 and γ_1 are $k \times 1$ vectors of the VECM's intercepts and linear trend coefficients respectively.

In line with Bierens and Martins (2010), for the $k \times 1$ sequence Z_t, we assume that for some t there are fixed $r < k$ linearly independent columns of the time-varying $k \times r$

[15] Cheung and Lai (1993) have documented that the trace test shows more robustness to both skewness and excess kurtosis in the residuals than the maximal eigenvalue test.

matrix $\beta_t = (\beta_{1t}, \ldots, \beta_{rt})$, that forms the time-varying space of the cointegrating vectors. We can write the time-varying Vector Error Correction model (TV-VECM) of order p as:

$$\Delta Z_t = \gamma_0 + \alpha\beta_t' Z_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta Z_{t-j} + \varepsilon_t, \qquad t = 1 \ldots T \tag{3}$$

where $\varepsilon_t \sim N_k(0, \Omega)$, α is a fixed $k \times r$ matrix with rank r, β_t is a time-varying $k \times r$ matrix also with rank r, T is the number of observations, Ω and Γ_j are $k \times k$ matrices. We test the null hypothesis of time-invariant (TI) cointegration Π_t'=Π'=$\alpha\beta'$, against the alternative time-varying (TV) cointegration of the type $\Pi_t = \alpha\beta_t'$.

Bierens and Martins (2010) (Lemma 1) prove that under standard smoothness and orthonormality conditions, the parameters of the TV cointegrating vector β_t can be approximated for some fixed m by a finite sum of Chebyshev time polynomials $P_{i,T}(t)$ of decreasing smoothness:

$$\beta_t = \beta_m(t/T) = \sum_{i=0}^{m} \xi_{i,T} P_{i,T}(t), \quad t = 1 \ldots T, \tag{4}$$

where $1 \leq m < T - 1$ and $\xi_{i,T} = \frac{1}{T}\sum_{t=1}^{T} \beta_t P_{i,T}(t)$ for $i = 0, \ldots, T$-1, are unknown $k \times r$ matrices. Chebyshev time polynomials are defined by: $P_{0,T}(t)$= 1 and $P_{1,T}(t)$= $\sqrt{2}\cos\left(\frac{i\pi(t-0.5)}{T}\right)$, $t = 1, 2, \ldots T$, $i \geq 1$, such that, for all couples of integers i, j, the following orthonormality property holds: $\frac{1}{T}\sum_{t=1}^{T} P_{i,T}(t) P_{j,T} = 1(i = j)$.

Testing for TV cointegration corresponds to the null and alternative hypothesis:
For TI: $H_0 : \ \xi_{i,T} = O_{k\times r}$ for i=1,$\ldots, m$, and $\xi_i = O_{k\times r}$ for $i > m$.
For TV: $H_1 : \ lim_{T\to\infty} \xi_{i,T} \neq O_{k\times r}$ for some $i = 1, \ldots, m$, and $\xi_i = O_{k\times r}$ for $i > m$.

Substituting (4) in (3), we get: $\Delta Z_t = \gamma_0 + \alpha \left(\sum_{i=0}^{m} \xi_i P_{i,T}(t)\right)' Z_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta Z_{t-j} + \varepsilon_t$, which can be rewritten as:

$$\Delta Z_t = \gamma_0 + \alpha\xi' Z_{t-1}^{(m)} + \sum_{j=1}^{p-1} \Gamma_j \Delta Z_{t-j} + \varepsilon_t, \tag{5}$$

where $\xi' = \left(\xi_0', \xi_1', \ldots, \xi_m'\right)$ and $Z_{t-1}^{(m)} = P_{i,T}(t) \otimes Z_{t-1}$, for $i = 0, \ldots, m$.

The null hypothesis of TI cointegration corresponds to ξ'=$(\beta, O_{r,k.m})'$ so that $\xi' Z_{t-1}^{(m)}$= $\beta' Z_{t-1}$. We can estimate both (5) and its time invariant counterpart (2), which is equivalent to (5) with $m = 0$, by maximum likelihood. The likelihood ratio test for the null hypothesis of standard (TI) cointegration against the alternative TV cointegration is given by:

$$LR_T^{tvc} = -2\left[\hat{l}_T(r, m=0) - \hat{l}_T(r, m)\right] \tag{6}$$

where $\hat{l}_T(r, \cdot)$ are the log-likelihoods computed in the estimated values of the VECM parameters. When $m = 0$, we are in a TI case, when $m > 0$, we are in a TV case. In both cases, r is the cointegration rank which we assume to be given. Bierens and Martins (2010) prove in Theorem 1 that given m, $r \geq 1$, under the null hypothesis of standard cointegration, the LR_T^{tvc} statistic defined above is asymptotically distributed as a chi-squared distribution with $r \times m \times k$ degrees of freedom.[16]

4. Empirical results

4.1 The Real Price of Gold and Silver

Jastram (1978) pointed out that the purchasing power of gold depends on the relation of commodity prices to gold prices. The author used the term "Retrieval Phenomenon" to describe the cyclical swings of the commodity price level around the level of gold. Due to this phenomenon, gold maintains its purchasing power over the long-run.[17]

Another way to assess how effective gold and silver have been as inflation hedges is to examine the historical fluctuations in the real (inflation-adjusted) prices (Erb and Harvey, 2013). If gold (silver) was a perfect inflation hedge, the real price of gold (silver) would be stationary. In that respect, we employ four alternative unit root tests to examine the stationarity properties of the real price of gold and silver: (i) Dickey and Fuller (1979, henceforth ADF), (ii) Phillips and Perron (1988, henceforth PP), (iii) the Zivot and Andrews (1992) unit root test with one break and (iv) the Lee and Strazicich (2003) for two breaks. The two break unit root test of Lee and Strazicich (2003) has been employed in order to counterbalance the potential loss of power of tests that ignore more than one break. Table 3 reports the results of the stationarity tests applied to the real price of gold and silver for the UK and the US. There is strong evidence at the 5% level that the real price of gold and silver is stationary. This conclusion is supported by both the Zivot and Andrews (1992), and Lee and Strazicich (2003) unit root test that

[16] Bierens and Martins (2009) compute empirical critical values via Monte Carlo simulations and show that even for T = 100 these values are very close to the asymptotic critical values.

[17] Jastram (1978) used the term "golden constant" to communicate his belief that the real price of gold maintains its purchasing power over long periods and that gold's long-run average real return has been zero. Nonetheless, for Jastram (1978), the short-run was a period of few years, while the long-run refers to many decades.

allow for structural break(s).

Next, it would be interesting to examine how these variables evolved over the 200 year period, under different inflation regimes.[18] Figures 4 and 5 illustrate the real price of gold and silver on US and UK local currencies since 1791 (1792 for silver). The price of both metals is deflated by the CPI series. In Figures 4 and 5 the darker shaded areas denote the inflationary periods, while the brighter shaded areas denote the deflationary periods.

The evidence drawn from the US experience (Figure 4) reveal that in five out of six major inflationary periods of US history since the eighteenth century, gold has lost its purchasing power while it has increased in two out of three deflationary periods. After 1971 this phenomenon has reversed, with gold to gain power in two inflationary periods and looses power in the Great Moderation period. For the US investor, silver holds its purchasing power in two out of three deflationary periods while losses its power in inflationary periods again until around 1930. In 1933 President Roosevelt took the US off the gold standard. Thereafter, the purchasing power of silver increased in inflationary periods and decreased in deflationary periods.

The UK evidence (Figure 5) demonstrates that before 1971 when the price of gold was freed, gold losses its operational power in all three inflationary periods while holds or even increases its operational power in the three deflationary periods. In the period after 1971, this phenomenon was altered, with gold to gain power in the two inflationary periods and looses power during the Great moderation period. Silver holds its purchasing power in two out of three deflationary periods while losses its power in inflationary periods until around 1930. After the decision of the British government to suspend the gold standard, the operational wealth of silver increased in inflationary periods and decreased in deflationary periods until recently.

4.2. Time-invariant (TI) and time-varying (TV) cointegration analysis

Having established that the nominal gold, silver prices and the alternative inflation measures are all $I(1)$ (see Table 2), we proceed to the bivariate cointegration analysis. Since cointegration tests are sensitive to the lag length, we estimate a VECM with up to 12 lags in each case and use both the Akaike (AIC) and the Schwarz (SIC)

[18]We have adopted the inflationary and deflationary periods proposed by Jastram and Leyland (2009). The authors defined those periods in a descriptive sense of price behavior. Inflation refers to a period of rapidly rising prices while deflation connotes an interval of swiftly falling prices. Although, the 1980-2000 period could be considered as part of Great Moderation period rather than a deflationary period, while the 2001-2010 period could be better described as 'inflation uncertainty' period rather than 'inflationary period'.

information criteria for additional robustness.[19] Another important issue is whether the deterministic component is present in the long-run relationship. Since both the price of precious metals and the consumer prices may include a trend, we follow Doornik et al. (1998) notation and adopt two trend specifications for each model: the Case I model (H_c) with a constant to lie in the cointegration space and the Case II model (H_l) with unrestricted constant and restricted trend that allows for non-zero drift in any unit-root processes found in the cointegration analysis.

In the generalized Fisher effect framework, our definition of long-run inflation hedging is derived from the comovement of gold or silver nominal price with the consumers prices in the following time-series regression:

$$PM_t = a + \beta CPI_t + u_t \tag{7}$$

where PM_t and CPI_t denote the natural logarithm of precious metals (gold or silver) prices and consumers prices respectively and u_t is the error. The coefficient β is the LR elasticity of gold and silver prices with respect to goods prices, indicating the percentage change in precious metals prices for every 1% change in goods prices. Possible outcomes once the null of no cointegration is rejected, include $0 < \beta < 1$ (partial hedge), $\beta = 1$ (full hedge), and $\beta > 1$ (superior performance).

The results of the Johansen (1995) (TI-VECM) and the Bierens and Martins (2010) (TV-VECM) cointegration tests for the US and the UK are summarized in Table 4. The pairs of variables which did not show any evidence of time-invariant or time-varying cointegration were excluded from the subsequent analysis.[20]

4.2.1. TI cointegration analysis

The results for the Johansen's trace test determine whether a long-term relation exists between each pair of gold (silver) prices and consumer prices. The null hypothesis is that there is no cointegrating relation, and if this is rejected, we test the hypothesis that there is at most one cointegrating vector. Results in Table 4 suggest that there is one cointegrating vector between each pair of gold prices and headline inflation in both countries according to the AIC (weaker evidence was found for SIC). The evidence for silver prices and headline inflation indicates no long-run relationship for both countries. We also observe strong evidence for a cointegrating relationship for

[19]There are cases where the two information criteria indicate the same lag length. The results in these cases are identical.

[20]The numerical time-invariant and time-varying cointegration test results are available from the authors upon request.

gold prices and expected CPI measures (HP, CF) for the UK. The same holds for the US (AIC), though the evidence is weaker for the SIC. A weak long-run relationship between silver prices and CF is also evident (AIC) for the UK and in Case II models for the US. Strong long-run relationship between gold price and the three core price measures is found for Case I model in the UK (moderate evidence was found for Case II model). Similar result emerges for the US (AIC). The SIC criterion supports cointegration in fewer cases for both countries. No significant cointegrating relation is present between silver prices and core price level measures.

Table 5 reports the long-run estimates for the relationship between gold and the alternative consumer price measures for the UK and the US. The Johansen trace test has not shown any significant long-run relationship between silver prices and the alternative CPI measures. The cointegration vectors have been normalized on gold and in each case the CPI coefficients have the expected sign. Turning to the magnitudes of the long-run coefficients, the estimated point coefficients for the US (Panel A in Table 5) range from 1.24 (HD) to 1.61 (CF). As can be seen most of the CPI coefficients are highly significant, indicating a positive relationship between US gold prices and the various consumer price measures. For the UK (Panel B in Table 5) the estimated point coefficients range between 1.03(*ses*) and 1.33(*d*4). It is also evident from the results that the long-run coefficient for the expected price level is higher compared to that of headline and core price level for the US. The highest coefficient for the UK is also observed for the expected CPI measures (CF) for most cases. Overall, gold provide superior hedge of future inflation in the long-run and this ability is stronger in the US (average *LR* 1.36) compared to the UK (average *LR* 1.16).

4.2.2. TV cointegration analysis

Over the last two centuries, the UK and the US have witnessed policy regime shifts and changes in market conditions. These events could affect the long-run relationship between precious metals and consumer prices. Furthermore, previous evidence of non-linearity for gold and silver (Frank and Stengos, 1989) reinforces the argument for the time-varying approach. For this purpose, we relax the assumption that the long-run relationship has remained constant through the last two century period by employing the time-varying framework of Bierens and Martins (2010), where the cointegrating vectors fluctuate over time and the Johansen set up is considered as a special case of the model.[21] Evidence of time-varying cointegration at each significance level will be

[21]The asymptotic p-values of Bierens and Martins (2010) test, for different combinations of the order m of the Chebyshev polynomial expansion and the lag order p, are presented in Bierens and Martins (2009).

referred as the case where the null hypothesis is rejected for the Chebyshev polynomial expansion up to order four ($m = 1 \ldots 4$). For example, if the null hypothesis for Chebyshev polynomial of order one ($m = 1$) is rejected at the 10% and the rest three ($m = 2 \ldots 4$) at the 5% level, this will be referred as rejection at the 10% level. If at least in one m we cannot reject the null, then no time-varying cointegration is considered.[22]

For the pair of US gold price and headline CPI (Panel A in Table 4), strong evidence emerges of a time-varying cointegration at the 5% level in both models (AIC) and in Case I model (SIC). The evidence for the Case II model is at the 10% level. Results for the UK (Panel B in Table 4) are in favor of a time-varying long-run relation between gold prices and headline CPI in Case I model (AIC and SIC) at the 5% and 10% significance level respectively. For the pair of silver prices and headline CPI, a time-varying long-run relation was found at the 5% level.

A significant time-varying long-run relation is detected between US gold prices and all the expected CPI measures at the 5% level (AIC) while according to SIC a significant relationship at the 5% was found against CF in both models (Panel A, Table 4). For the pair of US silver prices and expected measures, significant time-varying relationship was found in all models for the CF measure. A significant long-run relation is found between UK gold prices and the CF at the 10% level across all models. For the HP, there is evidence in favor of time-varying cointegration (AIC). For the UK silver prices, there exists a time-varying cointegration relationship with the expected price measures (HP, CF) according to both information criteria at the 5% level.

The time-varying cointegration test between US precious metal prices and core price measures reveals a significant time-varying relationship between US gold prices and the three core measures according to both information criteria. For the US silver prices and core measures, a weak time-varying relationship was detected in Case I models for *ses* at the 10% level. Evidence provided for UK shows a time-varying long-run relationship between UK gold prices and two core measures at the 5% significance level, namely $d4$ and $la8$, for both information criteria. Evidence of time-varying cointegration also found for the UK silver prices and all core price measures except the Case II model of $d4$ (both AIC and SIC).

In general, for the US there is evidence in favor of time-varying cointegration between gold prices and headline, expected and core measures. Similar results hold for the UK's gold prices while strong time-varying cointegration emerges between the UK's silver price and all price level measures. To sum up, we get stronger evidence of TV cointegration for both countries compared to the TI case. In the UK, we observe a

[22]Following Bierens and Martins (2010) since p-values are zero for any m larger than four we present the results until fourth order for the Chebyshev polynomial expansion.

wider comovement of precious metals with the alternative CPI measures (more cointegration pairs) whereas in the US the magnitude of the relationship is stronger within the TI framework with higher *LR* betas.

Since strong evidence is found in favor of time-varying cointegration, we present the plots of the time-varying coefficients for the cases that significant time-varying relationship has been detected at the 5% significance level. The time-varying coefficients β_{1t} and β_{2t} correspond to the cointegrating relationship $\beta_t^{'}Z_t = \beta_{1t}PM_t + \beta_{2t}CPI_t$ or $\beta_t^{'}Z_t = e_t$ where the process e_t represents the short-run deviations from equilibrium (real shocks or change in monetary regime).[23] Then β_t's will be approximated by $\beta_t(m) = \sum_{i=0}^{m} \xi_i P_{i,T}(t)$, where the ξ_i's indicate the Fourier coefficient.[24]

The plots of the CPI betas (β2t) for the US and the UK are presented in Figure 6. To conserve space, we plot only the headline (HD) and expected (HP) CPI betas. In the US (upper panel), we observe time variation in the gold-headline CPI (g_HD) relation. The headline inflation beta moves above its mean value (2.41) after the early 2000 period, a finding in line with Batten et al. (2014), who also report an increase in comovement between the two variables since 2002. Increased volatility is observed for the gold-HP beta up to the 1930s. The most recent decades, the HP betas move close to the average time-varying beta value (1.91), indicating a superior performance. In the UK (lower panel), a decrease in long term comovement between gold and HD is observed from the late 1990s until 2008. For the gold and HP pair, we observe relative instability up to the beginning of the 20th century and stability after. In the recent period the gold-HP beta moves towards zero.

In general, we observe instability in the time-varying beta coefficients. The time-varying betas of gold are on average above unity for both countries, providing a superior hedging ability for gold. The long-run elasticity of gold is more stable against expected inflation (HP) after 1930. The average time-varying beta coefficients of gold against the alternative CPI measures are lower for the UK compared to the US, confirming the evidence found in the time-invariant case.

[23]The Chebyshev polynomial ($P_{i,T}(t)$) order m and the VECM order p have been chosen according to the AIC. We thank Luis F. Martins for making the GAUSS code available.

[24]The components of time varying beta coefficients are rescaled to a maximum absolute value across variables and time of one. As Martins and Gabriel (2013) indicate, imposing the usual normalization in order to achieve linear dependence has consequences for time consistency of cointegration spaces. The authors conclude that this anomaly should be considered if there is evidence of structural breaks in the cointegrating relation.

5. Conclusion

This study has examined the long-run hedging properties of gold and silver against alternative measures of consumers price level for the UK and the US. A sample that spans for more than two centuries is employed. We consider a time-invariant and a time-varying approach in which the long-run cointegrating relation varies smoothly over time, allowing for non-linear adjustment in the long term dynamics.

First, we assess the two precious metal hedging effectiveness by examining the historical fluctuations in their real prices. If the real price of gold (silver) is mean reverting, then it follows that the series will return to its mean value. We show that the real price of gold and silver is stationary in both countries, after accounting for structural break(s).

Second, we found evidence that gold performance is superior against headline, expected and core CPI in the long-run. Within the time-invariant approach, the inflation hedging ability of gold is found on average higher in the US compared to the UK. In particular, the average gold's *LR* beta was found 1.36 for the former and 1.16 for the later. This finding is further confirmed by the time-varying approach. For gold, the time-varying betas against headline and expected inflation are higher than unity, affirming gold's superior hedging ability in both countries. Moreover, we display that the long-run relationship of gold against expected inflation became more stable over the last decades.

Finally, within the time-invariant approach, there is no long-run relationship between silver and all the price measures. The time-varying approach, where the adjustment is allowed to be nonlinear, overthrows this result: we found that there is strong evidence for time-varying long-run relation between silver prices and the alternative inflation measures for the UK while weaker evidence (if any) was found for the US. Overall, for both countries, there is stronger evidence in favor of time-varying long-run relationship between gold and silver prices and the alternative price level measures, relative to the time-invariant approach.

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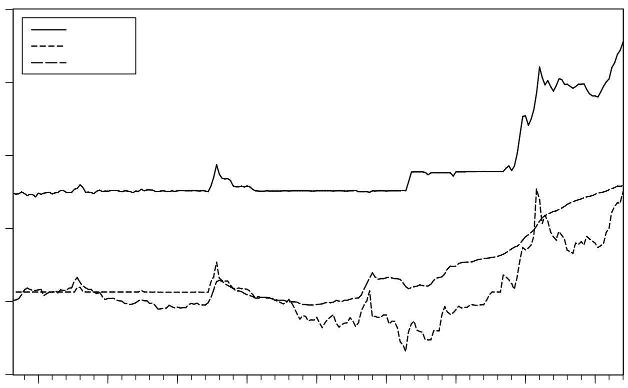

Fig. 1. US gold prices, silver prices and consumer price level (in logarithms)

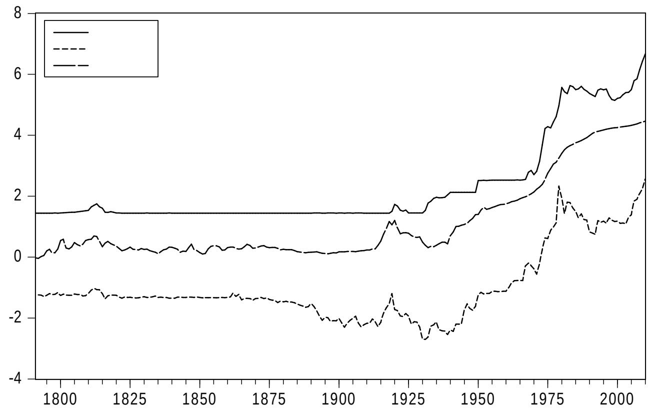

Fig. 2. UK gold prices, silver prices and consumer price level (in logarithms)

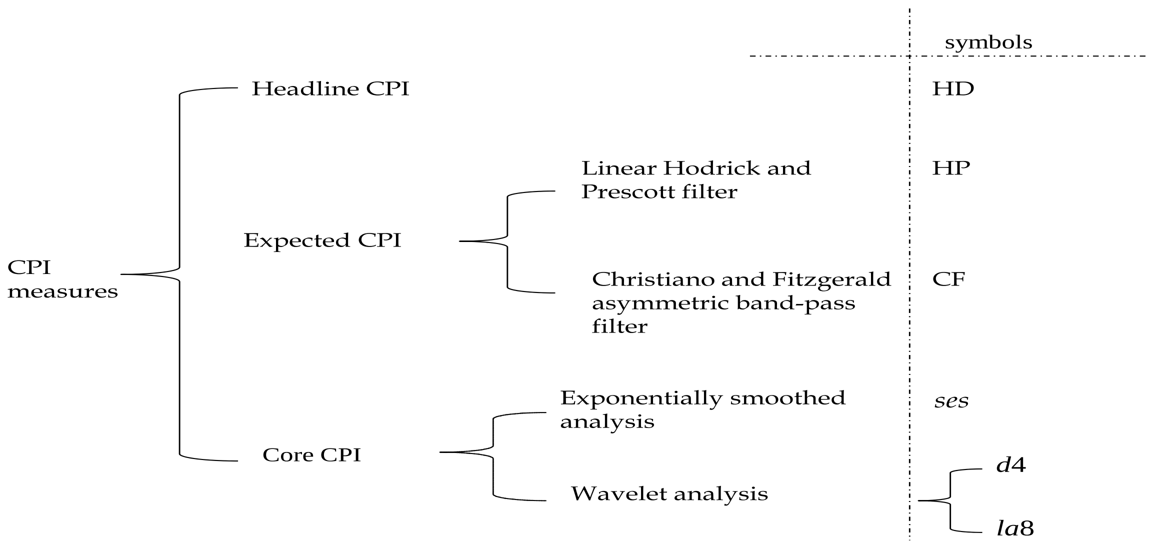

Fig. 3. Alternative inflation measures

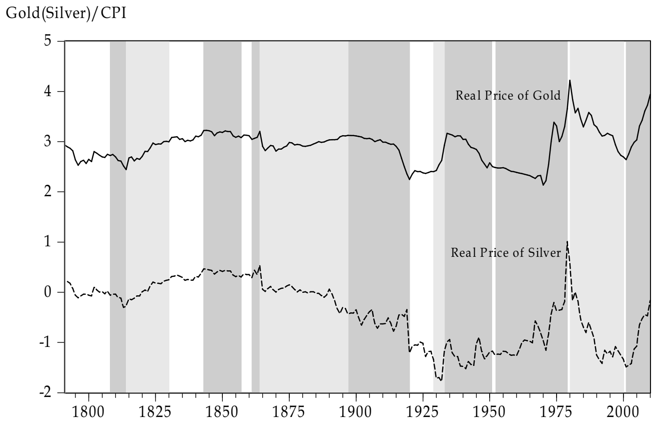

Fig. 4. The Real Price of Gold and Silver in US, 1791-2012

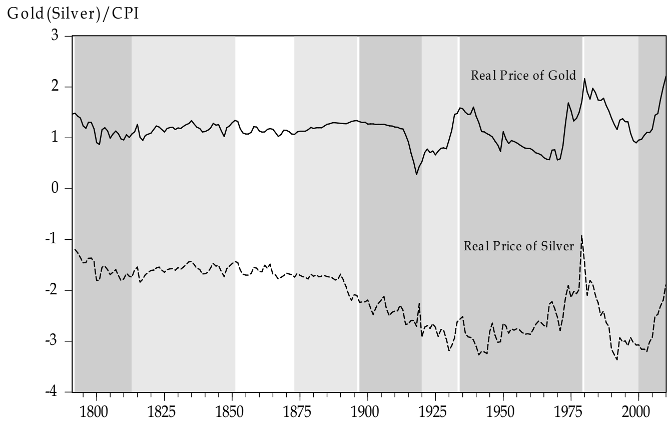

Fig. 5. The Real Price of Gold and Silver in UK, 1791-2012

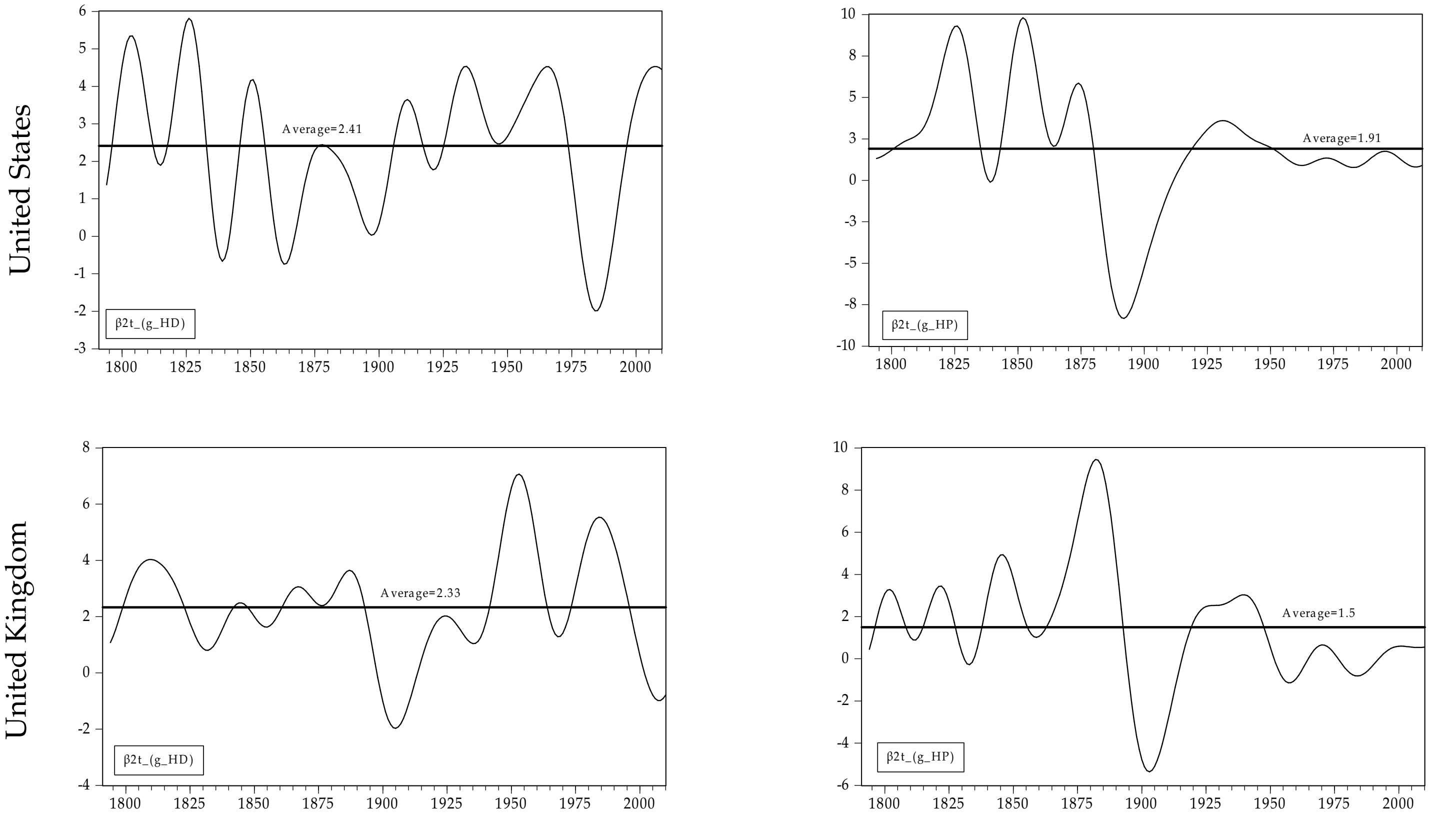

Fig. 6. US (upper panel) and UK (lower panel) time-varying cointegrating Vectors

Table 2. Unit root and stationarity tests

	US						UK					
	ADF		PP		NP		ADF		PP		NP	
Series	c	c,t	c	c,t	c	c,t	c	c,t	c	c,t	c	c,t
Log levels												
gold	0.99	0.99	0.99	0.99	4.60	0.08	0.99	0.99	0.99	0.99	4.33	0.61
silver	0.96	0.98	0.97	0.98	0.51	-1.79	0.99	0.99	0.99	0.99	2.81	-0.33
HD	0.99	0.97	0.99	0.99	3.20	-0.09	0.99	0.98	0.99	0.99	3.07	-0.03
HP	0.99	0.97	0.99	0.99	3.22	-0.09	0.99	0.97	0.99	0.99	3.01	-0.19
CF	0.99	0.10	0.99	0.99	3.21	-0.11	0.33	0.12	0.99	0.99	3.01	-0.16
*d*4	0.98	0.92	0.94	0.74	0.38	-2.48	0.99	0.96	0.92	0.76	0.37	-2.77
*la*8	0.85	0.61	0.94	0.64	0.18	-2.42	0.73	0.53	0.92	0.67	0.10	-2.70
ses	0.99	0.98	0.99	0.99	3.43	0.22	0.99	0.98	0.99	0.98	3.20	0.22
Returns (First differences)												
gold	0.00	0.00	0.00	0.00	-90.3***	-79.5***	0.14	0.00	0.00	0.00	-81.66***	-88.29***
silver	0.00	0.00	0.00	0.00	-90.3***	-78.7***	0.00	0.00	0.00	0.00	-107.86***	-98.64***
HD	0.00	0.00	0.00	0.00	-79.6***	-72.8***	0.00	0.00	0.00	0.00	-80.96***	-96.29***
HP	0.02	0.01	0.00	0.00	-13.0**	-17.9***	0.00	0.00	0.01	0.02	-19.94***	-21.17***
CF	0.72	0.00	0.01	0.03	-19.0***	-19.1***	0.76	0.00	0.02	0.05	-19.27***	-20.43***
*d*4	0.00	0.00	0.00	0.00	-81.4***	-36.7***	0.00	0.00	0.00	0.00	-54.63***	-28.79***
*la*8	0.04	0.13	0.00	0.00	-5.91*	-9.58**	0.14	0.41	0.00	0.00	-4.63	-6.40**
ses	0.00	0.00	0.00	0.00	-16.39***	-34.2***	0.00	0.00	0.00	0.00	-6.49**	-25.14***

Notes: The unit root tests are applied with (c,t) or without (c) a time trend. We report *p*-values for the Dickey and Fuller (1979) and the Phillips and Perron (1988) tests, while MZ_a is reported for the Ng and Perron's (2001). The lag length for the Dickey and Fuller (1979) test is selected via the Schwarz information criterion. The Phillips and Perron (1988) and the Ng and Perron's (2001) tests are based on the Bartlett kernel with bandwidth selected from the Newey–West method. ***, **, * denote rejection of the null-unit root hypothesis at the 1, 5 and 10% level respectively.

Table 3. Unit root tests for Real Price (RP) series

		ADF	PP	ZA		LS		
		t-stat.	t-stat.	t-stat.	TB		t-stat.	TB
RP_g_{us}	c	-3.12[6]**	-1.93	-4.67[6]**	1972	Model A	-3.80[8]*	1916, 1972
	c,t	-3.17[6]*	-2.03	-4.68[6]**	1972	Model C	-6.14[8]***	1835, 1953
RP_g_{uk}	c	-3.12[1]**	-1.88	-5.46[7]**	1972	Model A	-4.88[9]***	1915, 1979
	c,t	-3.18[1]*	-1.97	-5.5[7]**	1972	Model C	-6.15[11]***	1828, 1953
RP_s_{us}	c	-2.19[0]	-2.06	-3.25[2]**	1891	Model A	-4.09[13]**	1919, 1978
	c,t	-2.55[0]	-2.47	-3.7[2]**	1920	Model C	-5.30[13]**	1918, 1978
RP_s_{uk}	c	-2.46[0]	-2.31	-3.69[6]**	1891	Model A	-4.11[11]**	1919, 1978
	c,t	-2.91[0]	-2.61	-3.87[6]**	1920	Model C	-5.03[11]*	1919, 1978

Notes: ***, **, * indicate rejection of the null-unit root hypothesis at the 1, 5 and 10% level respectively. The number in the bracket shows the number of lagged difference terms in the corresponding unit root test. It was chosen by the Schwarz Criterion for ADF and ZA while for LS was chosen by the 't-sig' approach suggested by Perron (1997). We set an upper bound of fifteen for the lag length and test downwards until a significant (at the 10% level) lag is found. The Phillips and Perron (1988) test is based on the Bartlett kernel with bandwidth selected from the Newey–West method. In LS test, Model A allows for two shifts in the level, while Model C allows for two shifts in the level and the trend. TB denotes the time break date.

Table 4. Time-invariant (TI) and time-varying (TV) cointegration tests

	Case I: H_c^{AIC}		Case II: H_l^{AIC}		Case I: H_c^{SIC}		Case II: H_l^{SIC}	
	TI-VECM	TV-VECM	TI-VECM	TV-VECM	TI-VECM	TV-VECM	TI-VECM	TV-VECM
Panel A: United States								
Headline inflation								
(g_{us}, *HD*)	**	**	**	**	*	**	×	*
(s_{us}, *HD*)	×	*	×	×	×	*	×	×
Expected inflation measures								
(g_{us}, *HP*)	**	**	×	**	×	*	×	×
(g_{us}, *CF*)	**	**	**	**	**	**	**	**
(s_{us}, *HP*)	×	×	×	×	×	×	×	×
(s_{us}, *CF*)	×	**	*	**	×	**	*	**
Core inflation measures								
(g_{us}, *d*4)	×	**	×	**	×	**	×	**
(g_{us}, *la*8)	**	**	**	**	*	**	×	**
(g_{us}, *ses*)	**	**	**	**	**	**	×	×
(s_{us}, *d*4)	×	×	×	×	×	×	×	×
(s_{us}, *la*8)	×	×	×	×	×	×	×	×
(s_{us}, *ses*)	×	*	×	×	×	*	×	×
# of CI pairs	5/12	9/12	5/12	7/12	4/12	9/12	2/12	5/12
Panel B: United Kingdom								
Headline inflation								
(g_{uk}, *HD*)	**	**	**	×	×	*	*	×
(s_{uk}, *HD*)	×	**	×	**	×	**	×	**
Expected inflation measures								
(g_{uk}, *HP*)	**	**	**	*	**	×	×	×
(g_{uk}, *CF*)	**	*	**	*	**	*	**	*
(s_{uk}, *HP*)	×	**	×	**	×	**	×	**
(s_{uk}, *CF*)	×	**	*	**	×	**	*	**
Core inflation measures								
(g_{uk}, *d*4)	**	**	*	**	**	**	×	**
(g_{uk}, *la*8)	**	**	*	**	**	**	*	**
(g_{uk}, *ses*)	**	×	**	×	×	*	*	×
(s_{uk}, *d*4)	×	**	×	×	×	**	×	×
(s_{uk}, *la*8)	×	**	×	*	×	**	×	**
(s_{uk}, *ses*)	×	**	×	**	×	**	×	**
# of CI pairs	6/12	11/12	7/12	9/12	4/12	11/12	5/12	8/12

Notes: **, * indicate rejection of the null-unit root hypothesis at the 5% and 10% level of significance, respectively. TI-VECM and TV-VECM refer to the. Johansen (1995) and the Bierens and Martins (2010) cointegration tests, respectively. The critical values as well as the *p*-values of all the Johansen trace tests are obtained by computing the respective response surface according to Doornik (1998). $g_{us}(s_{us})$ denotes gold (silver) prices for the US. $g_{uk}(s_{uk})$ denotes gold (silver) prices for the UK. $H_c^{AIC}(H_l^{AIC})$ and H_c^{SIC} (H_l^{SIC}) denote the Case I (Case II) model, with the exponents AIC and SIC indicating the Akaike and Schwarz information criteria employed for lag length selection.

Table 5. Long-run relationship between gold prices and consumer prices

	Cointegrating vectors	Loading	Cointegrating vectors	Loading
Panel A: United States	H_c^{AIC}		H_l^{AIC}	
$(g_{us},\ HD)$	$g_{us} = 2.58^{***}_{(-28.42)} + 1.24^{***}_{(-12.26)} CPI_{HD}$	$\alpha_{g_{us},HD} = 0.004_{(0.42)}$	$g_{us} = -0.002_{(0,77)} t + 1.28^{***}_{(-9.27)} CPI_{HD}$	$\alpha_{g_{us},HD} = 0.012_{(1.31)}$
$(g_{us},\ HP)$	$g_{us} = 2.86^{***}_{(-30.5)} + 1.29^{***}_{(-12.3)} CPI_{HP}$	$\alpha_{g_{us},HP} = -0.0006_{(-1.01)}$		
$(g_{us},\ CF)$	$g_{us} = 2.8^{***}_{(-30.24)} + 1.37^{***}_{(-12.84)} CPI_{CF}$	$\alpha_{g_{us},CF} = 0.0002^{*}_{(1.77)}$	$g_{us} = -0.004^{**}_{(2.2)} t + 1.61^{***}_{(-11.17)} CPI_{CF}$	$\alpha_{g_{us},CF} = 0.00002^{***}_{(3.03)}$
$(g_{us},\ la8)$	$g_{us} = 2.85^{***}_{(-31.49)} + 1.27^{***}_{(-13.42)} CPI_{la8}$	$\alpha_{g_{us},la8} = 0.018_{(1.2)}$	$g_{us} = -0.002_{(0.96)} t + 1.35^{***}_{(-10.38)} CPI_{la8}$	$\alpha_{g_{us},la8} = 0.029^{*}_{(1.93)}$
$(g_{us},\ ses)$	$g_{us} = 2.85^{***}_{(-28.77)} + 1.3^{***}_{(-12.08)} CPI_{ses}$	$\alpha_{g_{us},ses} = 0.005_{(1.14)}$	$g_{us} = -0.001_{(0.5)} t + 1.32^{***}_{(-9.2)} CPI_{ses}$	$\alpha_{g_{us},ses} = 0.008^{*}_{(1.74)}$
	H_c^{SIC}		H_l^{SIC}	
$(g_{us},\ HD)$	$g_{us} = 2.89^{***}_{(-15.85)} + 1.39^{***}_{(-8.63)} CPI_{HD}$	$\alpha_{g_{us},HD} = -0.004_{(-0.69)}$		
$(g_{us},\ CF)$	$g_{us} = 2.8^{***}_{(-30.24)} + 1.37^{***}_{(-12.84)} CPI_{CF}$	$\alpha_{g_{us},CF} = 0.0002^{*}_{(1.77)}$	$g_{us} = -0.004^{**}_{(2.2)} t + 1.61^{***}_{(-11.17)} CPI_{CF}$	$\alpha_{g_{us},CF} = 0.00002^{***}_{(3.03)}$
$(g_{us},\ ses)$	$g_{us} = 2.82^{***}_{(-21.57)} + 1.3^{***}_{(-10.5)} CPI_{ses}$	$\alpha_{g_{us},ses} = 0.007^{*}_{(1.69)}$		
LR beta	*min* 1.24	*max* 1.61	*Average* 1.36	
Panel B: United Kingdom	H_c^{AIC}		H_l^{AIC}	
$(g_{uk},\ HD)$	$g_{uk} = 1.09^{***}_{(-16.19)} + 1.12^{***}_{(-25.05)} CPI_{HD}$	$\alpha_{g_{uk},HD} = 0.018_{(0.99)}$	$g_{uk} = -0.001_{(0.92)} t + 1.15^{***}_{(-18.61)} CPI_{HD}$	$\alpha_{g_{uk},HD} = 0.03^{*}_{(1.74)}$
$(g_{uk},\ HP)$	$g_{uk} = 1.09^{***}_{(-16.43)} + 1.15^{***}_{(-25.07)} CPI_{HP}$	$\alpha_{g_{uk},HP} = -0.0006_{(-0.52)}$	$g_{uk} = -0.001_{(1.11)} t + 1.19^{***}_{(-19.03)} CPI_{HP}$	$\alpha_{g_{uk},HP} = 0.0004_{(0.31)}$
$(g_{uk},\ CF)$	$g_{uk} = 1.08^{***}_{(-18.32)} + 1.17^{***}_{(-25.43)} CPI_{CF}$	$\alpha_{g_{uk},CF} = 0.00001_{(0.03)}$	$g_{uk} = -0.021^{*}_{(1.66)} t + 1.24^{***}_{(-20.95)} CPI_{CF}$	$\alpha_{g_{uk},CF} = 0.00001_{(0.91)}$
$(g_{uk},\ d4)$	$g_{uk} = 1.09^{***}_{(-15.44)} + 1.11^{***}_{(-23.49)} CPI_{d4}$	$\alpha_{g_{uk},d4} = 0.043_{(1.09)}$	$g_{uk} = -0.001_{(0.54)} t + 1.13^{***}_{(-16.96)} CPI_{d4}$	$\alpha_{g_{uk},d4} = 0.065_{(1.6)}$
$(g_{uk},\ la8)$	$g_{uk} = 1.09^{***}_{(-16.39)} + 1.16^{***}_{(-24.66)} CPI_{la8}$	$\alpha_{g_{uk},la8} = 0.01_{(0.38)}$	$g_{uk} = -0.001_{(0.37)} t + 1.17^{***}_{(-17.96)} CPI_{la8}$	$\alpha_{g_{uk},la8} = 0.024_{(0.84)}$
$(g_{uk},\ ses)$	$g_{uk} = 1.09^{***}_{(-16.73)} + 1.15^{***}_{(-25.23)} CPI_{ses}$	$\alpha_{g_{uk},ses} = 0.011_{(1.29)}$	$g_{uk} = -0.001_{(0.63)} t + 1.17^{***}_{(-18.79)} CPI_{ses}$	$\alpha_{g_{uk},ses} = 0.017^{\star}_{(1.92)}$
	H_c^{SIC}		H_l^{SIC}	
$(g_{uk},\ HD)$			$g_{uk} = -0.002_{(1.05)} t + 1.15^{***}_{(-13.06)} CPI_{HD}$	$\alpha_{g_{uk},HD} = 0.045^{**}_{(2.72)}$
$(g_{uk},\ HP)$	$g_{uk} = 1.09^{***}_{(-8.57)} + 1.24^{***}_{(-14.67)} CPI_{HP}$	$\alpha_{g_{uk},HP} = -0.0008^{*}_{(-1.63)}$		
$(g_{uk},\ CF)$	$g_{uk} = 1.08^{***}_{(-18.32)} + 1.17^{***}_{(-25.43)} CPI_{CF}$	$\alpha_{g_{uk},CF} = 0.00001_{(0.03)}$	$g_{uk} = -0.021^{*}_{(1.66)} t + 1.24^{***}_{(-20.95)} CPI_{CF}$	$\alpha_{g_{uk},CF} = 0.00001_{(0.91)}$
$(g_{uk},\ d4)$	$g_{uk} = 1.2^{***}_{(-6.19)} + 1.33^{***}_{(-11.18)} CPI_{d4}$	$\alpha_{g_{uk},d4} = -0.026_{(-1.44)}$		
$(g_{uk},\ la8)$	$g_{uk} = 1.09^{***}_{(-15.05)} + 1.14^{***}_{(-23.23)} CPI_{la8}$	$\alpha_{g_{uk},la8} = 0.013_{(0.46)}$	$g_{uk} = -0.0003_{(0.16)} t + 1.15^{***}_{(-16.28)} CPI_{la8}$	$\alpha_{g_{us},la8} = 0.025_{(0.87)}$
$(g_{uk},\ ses)$			$g_{uk} = -0.003_{(1.47)} t + 1.03^{***}_{(-13.91)} CPI_{ses}$	$\alpha_{g_{uk},ses} = 0.033^{***}_{(3.93)}$
LR beta	*min* 1.03	*max* 1.33	*Average* 1.16	

Notes: *,**,*** denote significance at the 10%, 5% and 1% level respectively. Numbers in parentheses are the values of the t-statistic. H_c^{AIC}(H_l^{AIC}) and H_c^{SIC} (H_l^{SIC}) denote the Case I (Case II) model, with the exponents AIC and SIC indicating the Akaike and Schwarz information criteria employed for lag length selection.

Appendix

Data description

Reinhart and Rogoff (2011) used as measures of inflation the consumer price indices or their close relative, cost-of-living indices. Since their analysis spans several earlier centuries, they rely on the meticulous work of a number of economic historians who have constructed such price indices item by item, most often by city rather than by country, from primary sources. When more than one city index is available for a country, they work with the simple average across cities (or regions) for the same country, such as in much of the pre-1800s data. [25]

The entire exchange-rate series is distinctive in two respects.[26] First, with rare exception (fourth quarter of 1833 and all of 1834), the data refer to actual and large-scale transactions rather than advertised, posted, or otherwise hypothetical exchange rates (the latter commonly recorded until the late nineteenth century). Second, the data are annual averages, covering as much of each year as possible, rather than pertaining to a specific day or month of the year. For 1870-1914, the data are annual averages of daily rates. For 1791-1869, the data are annual averages of quarterly values, these values derived as averages of all available intra-quarterly observations. For 1791-1912 the exchange-rate data pertain to what was called a "sight" (or "demand") bill of exchange. This meant that the buyer of British pounds paid in dollars immediately, but received the pounds after shipping the bill across the Atlantic and "presenting" it in London. Until 1879, in fact, "time" bills were the basis of exchange transactions. For example, a 60-day bill would involve an additional 63-day lag before receiving pounds-60 days inherent in the bill itself plus three "days of grace". The time-bill data (1791-1878) are converted to a sight-bill basis by eliminating the interest-component associated with the additional lag beyond that for a hypothetical sight bill. For 1913 and later years the exchange-rate data are for "cable transfers," whereby pounds are received on the same day that dollar payment is made. By the year 1913, the difference between the sight and cable rate is so small as to be unimportant for most purposes.[27] Officer (p. 13-19), provides a comprehensive description of the London market price of gold.[28]

[25] The complete references by author and period to this body of work are provided in Carmen Reinhart's webpage. Available at: http://www.carmenreinhart.com/data/browse-by-topic/topics/2/

[26] For further details on the exchange-rate series, the interested reader may consult Lawrence H. Officer, Between the Dollar-Sterling Gold Points (Cambridge University Press, 1996) and Lawrence H. Officer, "Exchange Rates," in Susan B. Carter, Scott S. Gartner, Michael Haines, Alan Olmstead, Richard Sutch, and Gavin Wright, eds. Historical Statistics of the United States: Millennial Edition, vol. 5 (Cambridge University Press, 2006).

[27] Source: Lawrence H. Officer, "Dollar-Pound Exchange Rate From 1791," MeasuringWorth, 2011 URL: http://www.measuringworth.com/exchangepound/

[28] Available at: http://www.measuringworth.com/docs/GoldBackground.pdf