Is Gold an Inflation-Hedge? Evidence from an Interrupted Markov- Switching Cointegration Model[$]

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Abstract

This paper investigates the inflation hedging role of gold price after controlling for the prices of other investment assets. We use annual data on the U.S. economy spanning from 1833 to 2013. We employ a recently developed flexible nonlinear approach that allows for potential 'interruption' in the long run equilibrium relationship in which the equilibrium term dynamics is modelled as an AR(1) depending upon an unobserved state process that is a stationary first-order Markov chain in two states, stationarity and non-stationarity. While, a battery of standard cointegration tests without and with breaks could not find evidence to support the inflation hedging role of gold, results from the flexible nonlinear approach indicate the existence of temporary cointegration between gold price and inflation during 1864, 1919, 1932, 1934, 1976, 1980 and 1982. The interruptions in the long-run relationship at different time periods seem to be associated with the different structural changes that affected the gold market.

Keywords: Gold; Inflation; Hedging; Interrupted Markov-Switching Cointegration.
JEL Codes: C22; C52; G15; Q02.

Introduction

Gold is usually considered an important asset by most investors, policy makers and academics due to a number of benefits accruing from it. This includes its role as an investment asset[1], as a

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[1] Gold is particularly considered an investment instrument or option for a number of reasons: It can easily be converted to cash (liquidity), maintains its value over time as a commodity largely due to its fixed quantity unlike fiat currency that can easily be printed by the authorities, an inflationary hedge, portfolio for diversification and hence risk reduction, its usefulness in the production of important products such as jewellery and electronics among others. There are a number of investment vehicles for gold. These include purchasing gold directly in form of Gold bars, bullion or coins, shares in gold mining companies, Derivatives such as gold forwards, futures and options, Gold accounts including unallocated and allocated as well as investment in an accumulation plan, Gold certificates

store of value, diversification benefit, financial arbitrage, political unrest, currency risk (i.e hedge against exchange-rate risk for investors with dollar holdings since gold is priced in dollars), potential safe haven or hedge against inflation (Chua and Woodward, 1982; Harmston, 1998; Worthington and Pahlavani, 2007; Baur and Lucey, 2010, Wang et al., 2011; Dee et al., 2013). Gold is always on high demand for jewellery, coins, bars and by many industries, such as electronics, space, as well as medical technology (Wang et al, 2011). Gold has characteristics that make it unique from other commodities: is durable, relatively transportable, universally acceptable and easily authenticated (Worthington and Pahlavani, 2007). Gold's positively skewed returns might also provide it safe-haven properties (Lucey, 2011). Although its role as a store of value in the monetary system diminished after the collapse of the Bretton Woods system in 1973, it is retained in many Asian countries and continues to hold a significant symbolic value even in countries where it plays lesser monetary role. These benefits notwithstanding, gold is being viewed by some as a "barbarous relic" with no modern role to play or "just another commodity" that rarely adds value in an investment strategy (Ranson and Wainright, 2005).

The main characteristics of gold are closely linked to their supply and demand factors. Generally, the supply of gold is relatively inelastic and stable owing to the tediousness of establishment of new mines, difficult extraction process, passive keeping of stocks of gold by Central Banks irrespective of the patterns of the real gold price (Aizenman and Inoue, 2012; Beckmann and Czudaj, 2013). However, the demand for gold is rapidly changing in response to global economic occurrences. Baur and McDermott (2010) separated the total demand for gold into three categories: demand for jewellery, demand for industrial and dental, and investment demand whereas Ghosh et al. (2004) simply divided it into two: the "use demand", where gold is used directly for producing jewellery, medals, coins, electrical components, among others and the "asset demand", where it is used by governments, fund managers and individuals as an investment to hedge inflation and other forms of uncertainty.

A key interest of long term investors is to protect their wealth against both expected and unexpected inflation; hence they are often concerned with maintaining the purchasing power of investment assets over time consistent with the Fisher (1896) hypothesis which was expanded by Fama and Schwert (1977). The aim of this study is to examine the hedging function of gold. When an asset co-moves with inflation, it can be viewed as an inflation hedge (Dee et al., 2013). Alternatively, a hedge is defined as an asset that is uncorrelated or negatively correlated with

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and Gold Exchange-traded products (ETPs) such as exchange-traded funds (ETFs), exchange-traded notes (ETNs), and closed-end funds (CEFs) which are traded like shares on the major stock exchanges (HSBC 2015; Smith 2015; Wikipedia, 2015).

another asset or portfolio on average while a safe haven is defined as an asset that is uncorrelated or negatively correlated with another asset or portfolio in times of market stress or turmoil (Baur and Lucey, 2010; Baur and McDermott, 2010). The gold return-inflation relationship reflects the extent to which gold is popular in the economy relative to the fiat money and other investment assets (Chua and Woodward, 1982). Wang, et al. (2011) also noted that the effectiveness of gold as a good hedge for both ex ante and post ante inflation depends on the economic conditions or characteristics of each country (Wang, et al., 2011). In this regard this study investigates whether gold price acts as a hedge to inflation for the case of USA.

The popular belief is that gold price tends to increase with the general price level hence providing a hedge against total inflation (Wang, et al., 2011; Dempster and Artigas, 2010). Theoretically, an increase in expected inflation will force investors to purchase gold, to either hedge against the expected decline in the value of money or speculate due to the associated rise of the price of gold. The resulting higher demand will lead to increasing gold price in time of the increasing inflation expectations (Beckmann and Czudaj, 2013). Therefore, knowledge regarding future inflation will enable investors to gain excess revenues by buying and selling gold in spot and futures markets in anticipation of prospective market adjustments. Consequently, the price of gold price would act as a leading indicator of inflation making gold an instrument for hedging against future inflation (Beckmann and Czudaj, 2013).

Empirically, a number of researchers have investigated the hedging or safe haven role of gold. However, the results are mixed and hence inconclusive. Mahdavi and Zhou (1997) test the performance of gold and commodity prices as leading indicators of inflation with cointegration and vector error correction model (VECM) using the Johansen framework for a quarterly sample period that ranges from 1970 to 1994 and conclude that consumer prices and the price of gold are not cointegrated. Their findings are consistent with those of Garner (1995) and Cecchetti, et al. (2000) who were also unable to empirically support the usefulness of the gold price as leading indicator for inflation. Adrangi et al. (2003) find that gold prices are positively correlated with expected inflation and conclude that a gold investment may be a reliable inflation hedge in both the short-run and the long-run. Levin and Wright (2006) analyze the short-run and long-run determinants of the gold price for the USA over a sample period from 1976 to 2005. They identify a stable long-run relationship between the gold price and the price level. Based on the traditional VECM they also provide evidence that the change in the gold price is positively related to the change in inflation, inflation volatility, and credit risk while its relationship with the U.S. dollar trade weighted exchange rate and the gold lease rate is negative.

In an empirical study based on monthly data covering September 1994 to December 2005 period for 14 countries: Australia, Canada, the European Union, New Zealand, Sweden, the United Kingdom, Japan, Mexico, Norway, the USA, Brazil, China, India, and Israel, Tkacz (2007) conclude that the gold price contains significant information for future inflation in several countries, especially those with formal inflation targets. Using a modified cointegration framework that allow for instabilities in the long run relationship for two sub-periods: 1945 to 2006 and 1973 to 2006, Worthington and Pahlavani (2007) find evidence in favour of a cointegrating relationship between the price of gold and inflation in both sample periods, thus concluding that gold can serve as an effective hedge for inflation. However, Blose (2010) using the U.S. data covering the period 1988-2008, finds no relationship between nominal gold returns and expected inflation.

Baur and Lucey (2010) study the relationship between U.S., U.K. and German stock and bond returns and gold returns to investigate gold as a hedge and a safe haven. They estimated regressions based on both full sample and subsample and find that gold is a hedge against stocks on average and a safe haven in extreme stock market conditions with the later property being short-lived. Wang et al. (2011) analyze the short-run and long-run inflation hedging effectiveness of gold in the USA and Japan using monthly data spanning from January 1971 to January 2010. They conduct the linear cointegration test proposed by Engle and Granger (1987) as well as the nonlinear threshold cointegration test suggested by Enders and Siklos (2001) and show that in low momentum regimes gold is unable to hedge against inflation in both the USA and Japan, however, in high momentum regimes, a gold investment is able to hedge against inflation in the USA, and partially hedge against inflation in Japan.

Dee et al. (2013) examine the hedging role of gold for stock and inflation in China mainland market. Using quantile regression and binary probit model, they find that gold cannot always hedge stock and inflation risk for short-term investors, while it is a good hedge for stock or inflation in the long term. However, they could not find evidence of its safe haven properties in the China market. Beckmann and Czudaj (2013) examine inflation hedging ability of gold using data for the USA, the UK, the Euro Area, and Japan covering from January 1970 to December 2011. They allow for nonlinearity by using Markov switching vector error correction model (MS-VECM) and also discriminate between long-run and time-varying short-run dynamics. They find that gold is partially able to hedge future inflation in the long-run with stronger ability for the USA and the UK compared to Japan and the Euro Area. They also conclude that the role of gold as an inflation hedge essentially depends on the time horizon and that one regime roughly corresponds to normal times while the other approximately accounts for turbulent times.

From the foregoing, it is obvious that the hedging function of gold is inconclusive. Results differ depending on methodology, sample period as well as country. Majority of the studies are based on the conventional cointegration tests and vector error correction models which assume that the parameters are constant over time. This assumption may be too restrictive as in reality gold price and inflation may fluctuate due to business cycles that may lead to nonlinearity in the relationship. The only studies that accounted for temporary or permanent, smooth or dramatic shifts in the gold price-inflation relationship are Wang et al. (2011) using threshold cointegration and Beckmann and Czudaj (2013) using Markov-switching VECM where the long-run vector switches across states, normal and turbulent times according to a Markov law.

The current study contributes to the debate about gold price hedging role in a number of ways. First, we use a newly developed flexible nonlinear and time-varying approach (a class of single-equation 'interrupted' cointegration model) by Martins and Gabriel (2014) which simultaneously accounts for long run co-movements amongst the variables, whilst endogenously allowing for well-documented structural shifts. In other words, the method we use allow for potential 'interruption' in the long run equilibria and it is based on break points generated stochastically and equilibrium term dynamics modelled as an AR(1) – depending upon an unobserved state process that is a stationary first-order Markov chain in two states, stationarity and non-stationarity. This is because changes in taste, technology, or economic policies shocks and herding behaviour can endanger long run stability and hence there may be cases in which cointegration does not hold for some periods of the sample while it holds for other periods. This notwithstanding, we also present results from other standard cointegration models. Second, we consider a very long sample (1833 - 2013) that enables us to capture as many potential structural shifts in the gold market and the U.S. economy as possible, including the recent financial and economic global crisis era. Third, we control for the prices of other investment assets (house price, silver price and stock price) while examining the inflation hedging ability of gold price, thereby addressing an important source of model misspecification typical of studies in this area, which generally just considers gold price and a measure of the price level.

The rest of the paper is organized as follows: the next section presents the econometric methodology used. Data is described in the third section. Section four presents the empirical results, while section five concludes.

Methodology

The econometric model employed is an interrupted cointegration single equation model. Although, modeling nonstationary cointegrated variables with globally stationary equilibrium errors, but locally nonstationary, is not completely new in the literature, this has not been fully explored. Hence an alternative flexible approach recently developed by Martins and Gabriel (2014) consistent with Psaradakis et al. (2004) is followed. Let $Y_t = \left(y_t, X_t{}' \right)'$ with $Y_t \in \Re^k$, $y_t \in \Re$ and $X_t \in \Re^{k-1}$. In its triangular representation, the model is defined as

$$y_t = \mu_t + \beta' X_t + u_t \tag{1}$$

where μ_t is the deterministic component (intercept, trend or time-dummies), β is a $(k-1)\times 1$ vector of coefficients, $X_t = X_{t-1} + v_t$, where v_t is a stationary zero-mean process that satisfies some functional central limit theorem, and

$$u_t = \rho_{s_t} u_{t-1} + \sigma_{s_t} \varepsilon_t \tag{2}$$

$\{s_t\}$ follows a stationary first-order Markov chain in $\{0,1\}$, with transition matrix, $P = (p_{ij})$, where

$$p_{ij} = P(s_t = j | s_{t-1} - i), \quad i, j \in \{0,1\} \tag{3}$$

and ε_t is assumed to be an $i.i.d\,(0,1)$ Gaussian distribution. It is assumed that $\{s_t\}$ is independent of $\{\varepsilon_t\}$. The autoregression (ρ) and variance (σ^2) parameters switch between states according to the unobserved Markov chain $\{s_t\}$. The existence of momentarily interruptions in the cointegration relationship requires the absence of absorbing states in the stationary and nonstationary parameter ranges $|\rho_0| < 1$ and $\rho_1 = 1$: $p_{ii} \in (0,1), i = 1,2.$

We obtain the corresponding error correction model (ECM) by rearranging terms

$$\Delta y_t = \mu_{\Delta t} + \alpha_{st} \begin{pmatrix} y_{t-1} - \mu_{t-1} - \beta' X_{t-1} \\ = u_{t-1} \end{pmatrix} + \eta' \Delta X_t + \sigma_{s_t} \varepsilon_t + u_t \tag{4}$$

where $\alpha_{st} = \rho_{st} - 1$, $\mu_{\Delta t} = \mu_t - \mu_{t-1}$, $\eta = \beta$. Here the standard cointegration regime $\alpha_0 < 0$ is interrupted in periods for which no error correction exists $\alpha_1 = 0$.

Consistent with Francq and Zakoïan (2001), the "equilibrium" error u_t can be characterized as a strictly and second-order (globally) stationary Markov-switching AR process, which do not require the usual local stationarity conditions (these are given by $|\rho_i| < 1, i = 0,1,$ for the second-order case). Let the 2×1 vector of ergodic probabilities ξ, that satisfy the condition $P\xi = \xi$, as $\xi_i = P(s_t = i), i \in \{0,1\}$, the unconditional probability that s_t will be in regime i at any given date. The sufficient conditions for the existence of a strictly and second-order stationary solutions of equation (2) are

$$\xi_0 \log|\rho_0| + \xi_1 \log|\rho_1| < 0 \text{ and} \tag{5}$$

$$(p_{00} + p_{10})\rho_0^2 + (p_{01} + p_{11})\rho_1^2 < 1 \tag{6}$$

respectively. Further, under the assumption of second-order stationarity and $|\rho_i| < 1, i = 0,1,$, the autocovariance function of $\{u_t\}$ is defined as

$$\mathrm{E}(u_t u_{t-h}) = l'Q^h V \tag{7}$$

where $l = (1,1)'$, $Q = \begin{pmatrix} p_{00}\rho_0 & p_{10}\rho_0 \\ p_{01}\rho_1 & p_{11}\rho_1 \end{pmatrix}$ and $V = \left(\xi_0 \dfrac{\sigma_0^2}{1-\rho_0^2}, \quad \xi_1 \dfrac{\sigma_1^2}{1-\rho_1^2} \right)'$

Interrupted cointegration is defined for $|\rho_0| < 1$ and $\rho_1 = 1$. Hence, $\{u_t\}$ is strictly stationary but finite moments will only exist (to obtain second-order stationarity) for

$$\rho_0^2 < \frac{(p_{00} - p_{11})}{(p_{00} - p_{11} + 1)} = 1 - \frac{1}{(p_{00} - p_{11} + 1)} \quad \text{with } p_{00} \geq p_{11} \tag{8}$$

In other words, interrupted cointegration in the weak sense, needs a "dominant" local stationary state, $s_t = 0$, that globally offsets the local nonstationarity of the equilibrium term.

The estimation of equations (1) and (2) is a two-stage procedure. The first step involves the estimation of (μ_t, β) by least squares and obtaining the residuals $\hat{u}_t$. Stock (1987) demonstrated that this estimator is superconsistent. The second step involves the use of maximum likelihood estimation, through numerical optimization or with an expectation maximization (EM) algorithm, to a two-regime Markov-Switching AR(1) model of the residuals $\hat{u}_t$, consistent with the procedures in Hamilton (1989), Hamilton (1994), and Kim and Nelson (1999).

The Markov-Switching model is specified as

$$\hat{u}_t = \rho_{s_t} \hat{u}_{t-1} + \sigma_{s_t} \varepsilon_t \tag{9}$$

with the set of unknown parameters

$$\theta = \{\rho_0, \rho_1, \sigma_0, \sigma_1, p_{11}, p_{22}\} \quad (10)$$

The estimation and inference of the parameters θ is carried out by maximizing the likelihood function $l(\theta)$ of the model. It involves recursive computation of probabilities about the unobserved regimes and obtaining $\hat{\theta}$ that maximizes the log-likelihood function.[2]

The interrupted cointegration assumes that $|\rho_0| < 1$ and $\rho_1 = 1$. Therefore, we test the standard null hypothesis, H_0: $\rho_1 = 1$. The likelihood ratio statistic

$$LR_T = 2(l(\hat{\theta}) - l(\hat{\theta}_R)) \quad (11)$$

is asymptotically $\chi^2_{(1)}$ -distributed, where $\theta_R = \{\rho_0, 1, \sigma_0, \sigma_1, p_{11}, p_{22}\}$

Data

We use annual data on gold and silver prices, consumer price index (CPI), house, and stock prices of the U.S. economy spanning from 1833 to 2013, with the start and end dates being purely driven by data availability on all the five variables under consideration. The data on nominal gold (Gold) and silver (Silver) expressed in U.S. dollar per ounce is obtained from www.kitco.com. Nominal SP500 stock price (Stock) and Winans International nominal house price (House) indexes are extracted from the Global Financial Database. The CPI is obtained from the website of Robert Sahr (http://oregonstate.edu/cla/polisci/sahr/sahr). All series are transformed into their natural logarithmic form. The variables have been plotted in Figure A1 in the Appendix.

The data are subjected to three unit root tests, Phillips and Perron (PP, 1988), Ng and Perron (NP, 2001), and Enders and Lee (EL, 2012) to determine their order of integration. While the PP and NP tests are standard unit root tests that do not account for breaks, the EL test allows us to accommodate for any number of structural breaks based on Fourier function. The results are presented in Table 1 for the case of intercept, and intercept and trend. The results show that all the series in levels: gold price, CPI, house price, silver price and stock price contain unit root as the tests cannot be rejected at any conventional level of significance. This implies that the series are non-stationary in levels. However, similar tests conducted on the first log differences shows that the series are all stationary as we are able to reject the null hypothesis of unit root at 1% level for all the series. Hence, we conclude that the series are integrated of order

[2] More technical details on the estimation and inferences can be found in Martins and Gabriel (2014).

one, I(1). This satisfies the condition for examining the long run relationship between the variables.[3]

Table 1: Unit root tests

	Phillips-Perron test			
	Level: Constant	Level: Constant & Trend	First-Differences: Constant	First-Differences: Constant & Trend
Gold	1.722	-0.440	-8.180**	-8.005**
CPI	1.550	-1.012	-6.055**	-6.108**
House	0.601	-2.376	-12.282**	-12.407**
Silver	0.210	-0.773	-11.927**	-12.151**
Stock	1.698	-1.338	-10.056**	-10.296**
	Ng-Perron test			
	Level: Constant	Level: Constant & Trend	First-Differences: Constant	First-Differences: Constant & Trend
Gold	3.575	-0.866	-11.588*	-115.136**
CPI	2.466	-2.637	-67.910**	-70.275**
House	1.980	-3.290	-88.925**	-89.054**
Silver	0.377	-2.283	-88.477**	-88.371**
Stock	2.379	-3.833	-84.251**	-124.984**
	Enders-Lee test			
	Level: Constant	Level: Constant & Trend	First-Differences: Constant	First-Differences: Constant & Trend
Gold	-2.597	-0.690	-5.055**	-9.165**
CPI	-0.059	-2.235	-16.633**	-10.946**

[3] The wavelet based unit root test of Gencay and Fan (2010) also confirmed that the series are I(1). Details of these results are available upon request from the authors.

House	-0.832	-3.814	-57.822**	-10.098**
Silver	-1.969	-2.818	-32.044**	-4.149**
Stock	-0.465	-3.339	-89.437**	-11.591**

** and * indicate significance at 1% and 5% level respectively.

Empirical Results

Prior to estimating the relevant models, we perform the Bai and Perron (2003) multiple break test (1 to M Globally determined breaks) and a trimming of 15%. The results are presented in Table 1. We find two significant break points (1933 and 1971) based on the powerful UDmax and WDmax statistics. This justifies the need for switching models.

Table 2: Multiple Structural Break Tests

UDmax statistic	981.250*
WDmax	1145.211*
No. of Breaks	2
Estimated Break dates	1933 1971

Note: * indicates significant at 5% level. UDmax critical values and WDmax critical values are 18.42 and 19.96 respectively.

However, we conducted a battery of conventional cointegration tests prior to implementing the interrupted cointegration which is the preferred method in this study. Most of the standard tests are single equation residual-based tests that allow for structural change. This is to enable us to consistently compare our results with those from the interrupted cointegration test. These results are presented in Table 3. The Hansen (1992a) test is both a test of parameter instability and cointegration test. In our case we reject at 1% significance level the null of cointegration for both the model with a level shift (C) and the model with level shift and trend (C/T). Also the null of cointegration is rejected at 1% for the Park (1992) added variable test. Therefore, we conclude based on Hansen's (1992a) and Park's (1992) tests, that gold price does not have a long-run relationship with price-level and other investment assets. Both Engle and Granger (1987) and Phillips and Ouliaris (1990) tests have the null hypothesis of no cointegration. We cannot reject this null hypothesis at any conventional level based on these two tests.

We also conduct the Gregory and Hansen (GH) (1996) test in which the cointegrating vector may be subject to a regime shift at an unknown time under the alternative hypothesis. This requires computing the test statistics (ADF, Zt and Zα) for all possible break points and selecting the smallest value obtained. The null of no cointegration for the GH test cannot be rejected for any of the test statistics. This is also robust to whether we assume that the structural change in the cointegrating relationship takes the form of a level shift (C), level shift with trend (C/T) or regime shift (i.e shift in the slope coefficient (C/S)). Overall, results from the standard cointegration tests do not provide evidence of cointegration between gold price and inflation.[4] This implies that the inflation hedging role of gold price is not supported by these tests.

Table 3: Conventional cointegration tests

Test	Statistic
Hansen Lc (C)	1.453**
Hansen Lc (C/T)	1.457**
Engle-Granger (C): tau	-2.617
Engle-Granger (C): z	-14.573
Engle-Granger (C/T): tau	-2.617
Engle-Granger (C/T): z	-14.573
Phillips-Ouliaris (C): tau	-2.878
Phillips-Ouliaris (C): z	-17.369

[4] In addition to the standard single equation residual-based cointegration tests, we also implemented the multivariate cointegration test by Johansen (1995). Both the Trace and Lambda-Max test statistics indicated no cointegration at 5% level of significance. However, when we implemented the time-varying version of this test proposed by Bierens and Martins (2010), there was strong evidence of cointegration, thus implying an inflation hedging role for gold prices – a result similar to that of Bampinasa and Panagiotidis (2015). Complete details of these results are available upon request from the authors. We, however, decided to work with the interrupted MS cointegration test, since unlike the time-varying cointegration, the former test allows us to detect cointegration at specific points in time, and hence, helps us to relate these periods of inflation hedging of gold prices with specific events in the economy. For the time-varying cointegration test, confidence sets needs to be created to see if gold is in fact serving as partial, full or superior hedge to inflation, depending upon whether the coefficient on price level is statistically less than, equal to and greater than one (Martins, 2015) - something Bampinasa and Panagiotidis (2015) did not perform, besides also ignoring possible misspecification of their results due to non-consideration of other assets in the portfolio of agents. In other words, for our purpose, the interrupted MS cointegration method is more informative than the time-varying cointegration approach.

Phillips-Ouliaris (C/T): tau	-2.878
Phillips-Ouliaris (C/T): z	-17.369
Park Added Variables	10.866**
GH-ADF(C)	-4.096
GH-Z_t(C)	-4.068
GH-Z_α(C)	-34.048
GH-ADF(C/T)	-5.199
GH-Z_t(C/T)	-5.072
GH-Z_α(C/T)	-45.420
GH-ADF(C/S)	-5.391
GH-Z_t(C/S)	-5.406
GH-Z_α(C/S)	-51.996

** and * indicate significance at 1% and 5% level respectively. Hansen's (1992a) and Park's (1992) tests have the null of cointegration, while the other tests have a null of no-cointegration. C stands for model with intercept only, C/T stands for model with intercept and trend, C/S stands for "regime shift" model where both the constant and slope parameters change.

The result for the Markov switching (MS) regression which accounts for the existence of short period interruptions in the cointegrating relationship between gold, inflation and the other investment assets is presented in Table 4 alongside the linear specification (no parameter switches). The results are obtained from the maximum likelihood estimation of $\hat{u}_t = \rho_{s_t}\hat{u}_{t-1} + \sigma_{s_t}\varepsilon_t$ (where $\hat{u}_t$ is the least square residuals) under the assumption that the unobserved process $\{s_t\}$ follows a stationary first-order Markov chain in $\{0,1\}$. It is important to note that the results presented in the lower panel of Table 4 relates to an MS model with constant transition probability where we have assumed that the Markov chain is not allowed to vary over time. In other words the MS model with constant transition probability assumes that the elements of the transition probability matrix (P_{ii}) are constant over time.[5] Our results show

[5] The transition probability is the probability that a Markov chain will move from one state (e.g. state 0) to the other state (e.g. state 1) (Piger 2007; Bazzi et al., 2014).

that the MS model has a better fit than the linear model as indicated by Akaike information criterion (AIC), Schwarz information criteria and Hannan-Quinn information criterion (HQC). Also the Hansen's (1992b) likelihood ration test rejects the null of a single regime at 1%. Hence, we turn to the results of the MS model specification.

Our results show that there is a change in the amount of variance of the error term as evidenced in the larger variance in regime 0 (σ_0) compared to that of regime 1 (σ_1) with both statistically significant at 1% level. Further, the autoregressive coefficient in regime 0 (ρ_0) is smaller than that of regime 1 (ρ_1). Again both coefficients are statistically significant. This may point to the possibility of a stochastic shift in the memory of the potentially equilibrium error which may capture shifts from stationarity/non-stationarity as long as $\rho_1 = 1.$ As it turns out, we observe that the null hypothesis $\rho_1 = 1$ cannot be rejected given that the *p*-value is greater than 10%. This implies that there is evidence of interrupted cointegration between gold price and inflation after controlling for other investment assets. We therefore conclude that gold acts as an inflation hedge.

The results based on an error correction specification of the interrupted cointegration as presented in lowest panel of Table 4 show that α_0 is negative and significant at 5% implying that any deviation in the long run equilibrium arising from short run shocks will be restored albeit slowly. Also the null hypothesis $\alpha_1 = 0$ cannot be rejected at any conventional level. In other words the standard cointegration regime $\alpha_0 < 0$ is interrupted in periods for which there is no error correction. Therefore, these results provide further evidence of interrupted cointegration.

We also date regime 1 (momentarily interrupted cointegration). Our results show that the probability of staying in a given state is high as the transition probabilities for regimes 0 and 1 are 0.80 and 0.95 respectively. This implies that the regimes are persistent. The average expected duration in regime 0 is about 5 years while it is about 21 years in regime 1. Diebold et al. (1994) and Filardo (1994) argue that the assumption of a constant transition probability matrix for a Markov switching model is too restrictive for many empirical settings. They extend the basic Markov switching model to allow the transition probabilities to vary over time using observable covariates.[6] Therefore we also estimate a model with time-varying transition probabilities, where

[6] The transition probability is time-varying if we assume that Markov chain is allowed to vary depending on the outcome of observed information. In other words the MS model with time-varying transition probability assumes that the elements of the transition probability matrix are allowed to be functions of past values of the dependent variable and of exogenous variables (Piger 2007; Bazzi et al., 2014).

the probability is a function of the last period's residual, i.e., u_{t-1}. We note that the fit of the MS model with constant transition probabilities is better than the MS model with time-varying transition probabilities based on the AIC, SIC and HQC.[7] However the average expected duration in regime 0 is same as in the constant model, while it is about double (43 years) for regime 1. Figure 1 compares the filtered regime probabilities (that is, the probability that the unobserved Markov chain is in a particular regime in period t, conditional on observing sample information up to period t) for the MS model with constant and time-varying transition probabilities associated with the local cointegration (regime 0). There is not much difference between the two regime probabilities except that the values for the constant transition probability model are slightly larger for some periods. From Figure 1, it can be observed that there was cointegration in 1864. There was however interrupted cointegration thereafter until 1918. There was evidence of cointegration momentarily in 1919, 1932, 1934, 1976, 1980 and 1982 with interruptions in between the various periods.[8] The interruptions in the cointegrating relationship between gold and inflation may be attributed to important periods of global structural changes such as the Black Friday in 1869 (the New York gold conspiracy), the Panic of 1873: a financial crisis that triggered a depression in Europe and North America and lasted from 1873 to 1879, the collapse of Bretton Woods which began with a temporary suspension of the dollar's convertibility into gold in 1971 and its final collapse in 1973, the oil price shocks in 1973 and 1979/80, the collapse of the USSR in 1991 and the subsequent democratization of Russia as one of the top producer of gold worldwide, the 1997 Asian financial crisis and its contagion, September 11, 2001 terrorist attack on the World Trade Center, the 'dot-com" bubble and burst in 2001 and the recent financial and economic crises initiated in 2007.

[7] Table A.1 in the Appendix reports the estimated parameters for the model with time-varying transition probabilities.

[8] We tried several samples: from 1833 to 1932, 1933 to 1971 and from 1972 to 2013 in line with the structural break tests and also 1970 to 2013 (as requested by a referee), but the dates of hedging was same as that of the full-sample. This is not surprising given that we are using a time varying approach which is capable of accounting for any break in the data. More importantly, the asset markets we consider are old markets, and have evolved over time, and this is what we aim to capture with our time-varying approach. Therefore, these results give us more confidence in the approach employed.

Table 4: Markov switching regression with constant transition probabilities

Parameter	Coefficient	Standard error
Linear Model:		
ρ	0.919	0.031
σ	0.119	
p-value (H_0: linear)	0.000	
AIC	-1.412	
SIC	-1.394	
HQC	-1.405	
Markov Switching Model:		
ρ_0	0.770**	0.163
ρ_1	0.962**	0.026
σ_0	-1.545**	0.151
σ_1	-2.572**	0.098
p_{00}	0.800	
p_{11}	0.953	
p-value (H_0: $\rho_1 = 1$)	0.147	
ECM:		
α_0	-0.141	0.067*
α_1	1.01E-04	2.86E-04
AIC	-1.641	
SIC	-1.535	

HQC	-1.598

** and * indicate significance at 1% and 5% level respectively

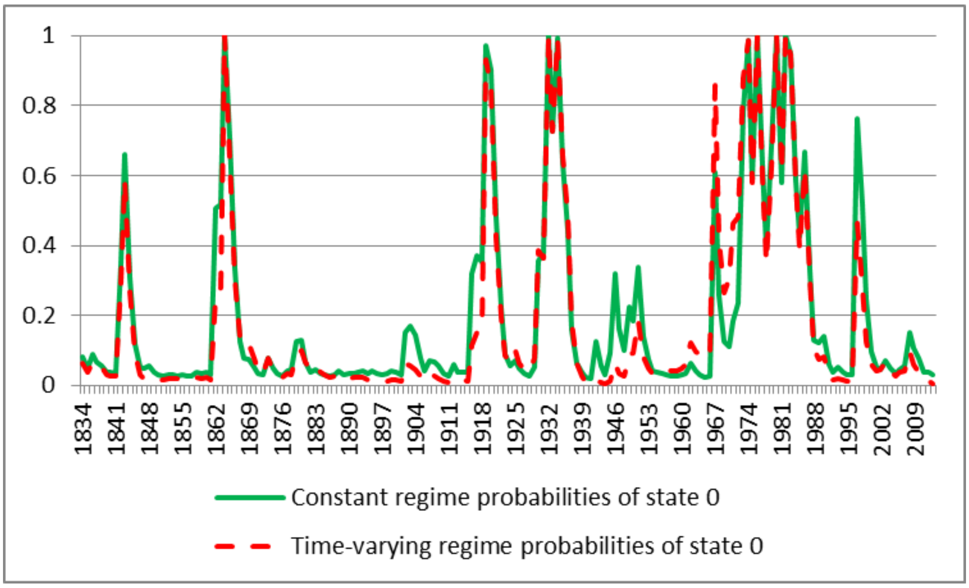

Figure 1: Interrupted cointegration in regime 0

Conclusion

There is a wide spread belief that gold acts as an inflation hedge. While theory appears to support this, the empirical evidences are mixed. This may be due to the differences in the methodology employed in analysing the relationship between gold and price-level. In this study, we examine the hedging role of gold against price-level, while controlling for other investment assets (silver, house and stock prices) under a nonlinear framework. We use data covering from 1833 to 2013. A Markov switching regression model that allows for interruption in the long run equilibrium relationship between gold price and inflation is used. However, prior to its implementation, we conduct a series of standard cointegration tests including that of Hansen (1992a) parameter instability test, Engle-Granger (1987), Phillips and Ouliaris (1990), Gregory and Hansen (1996), Park's (1992) added variables cointegration tests. Results from these conventional cointegration tests did not find evidence of cointegration between gold price and the CPI, controlling for other assets. However, when we confronted the same data with a more flexible method, we found evidence of cointegration and hence long run relationship between gold price and inflation. Particularly, there is evidence of interrupted cointegration which shows that the long run relationship is not constant at every point in time. This is also confirmed by results from the error correction specification. This implies that although gold may act as a hedge for inflation in the long- run, this role may be interrupted due to structural changes that affects the gold market. This finding has implications for academics, investors and policy makers. The changing nature of the comovement between gold and inflation implies that there cannot be a permanent panacea for exogenous shocks to the system. There may be need for short term measures such as temporary easing or tightening of monetary and fiscal policies and control of capital flows among others. However, it is also important to point out that these may then introduce policy uncertainty and consequently increase the cost of policy implementations. This finding will also help investors to make better asset allocation portfolio decisions. Although investors can protect their wealth by investing in gold, they need to monitor the changes in the market and economy in general to be able to effectively do so. Continuous and large hoarding of gold by investors and the general public as well as keeping of large reserves by Central Banks need to be cautiously done since the confidence may be eroded by structural changes. The findings of this study are also important for academics as the need to account for nonlinearities in modelling of gold price-inflation relationship has been demonstrated. As part of future research it would be interesting to extend our analysis to the short-run inflation hedging characteristic of gold as well.[9]

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[9] Using a DCC-GARCH model, we observed that gold returns and inflation are positively correlated for the entire sample barring the years of 1932 and 1933, i.e., during the Great Depression. So, in general, there is some preliminary evidence of gold serving as a hedge to inflation also in the short-run, but more analysis in this regard is required in the future. Complete details of these results are available upon request from the authors.

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Appendix

Table A.1: Markov switching regression estimates with time-varying transition probabilities:

Parameter	Coefficient	Standard error
ρ_0	0.802**	0.142
ρ_1	0.955**	0.028
σ_0	-1.504**	0.158
σ_1	-2.553**	0.086
p_{00}	0.798**	0.015
p_{11}	0.952**	0.047
p-value (H_0: $\rho_1 = 1$)	0.107	
AIC	-1.628	
SIC	-1.486	
HQC	-1.570	

** and * indicate significance at 1% and 5% level respectively.

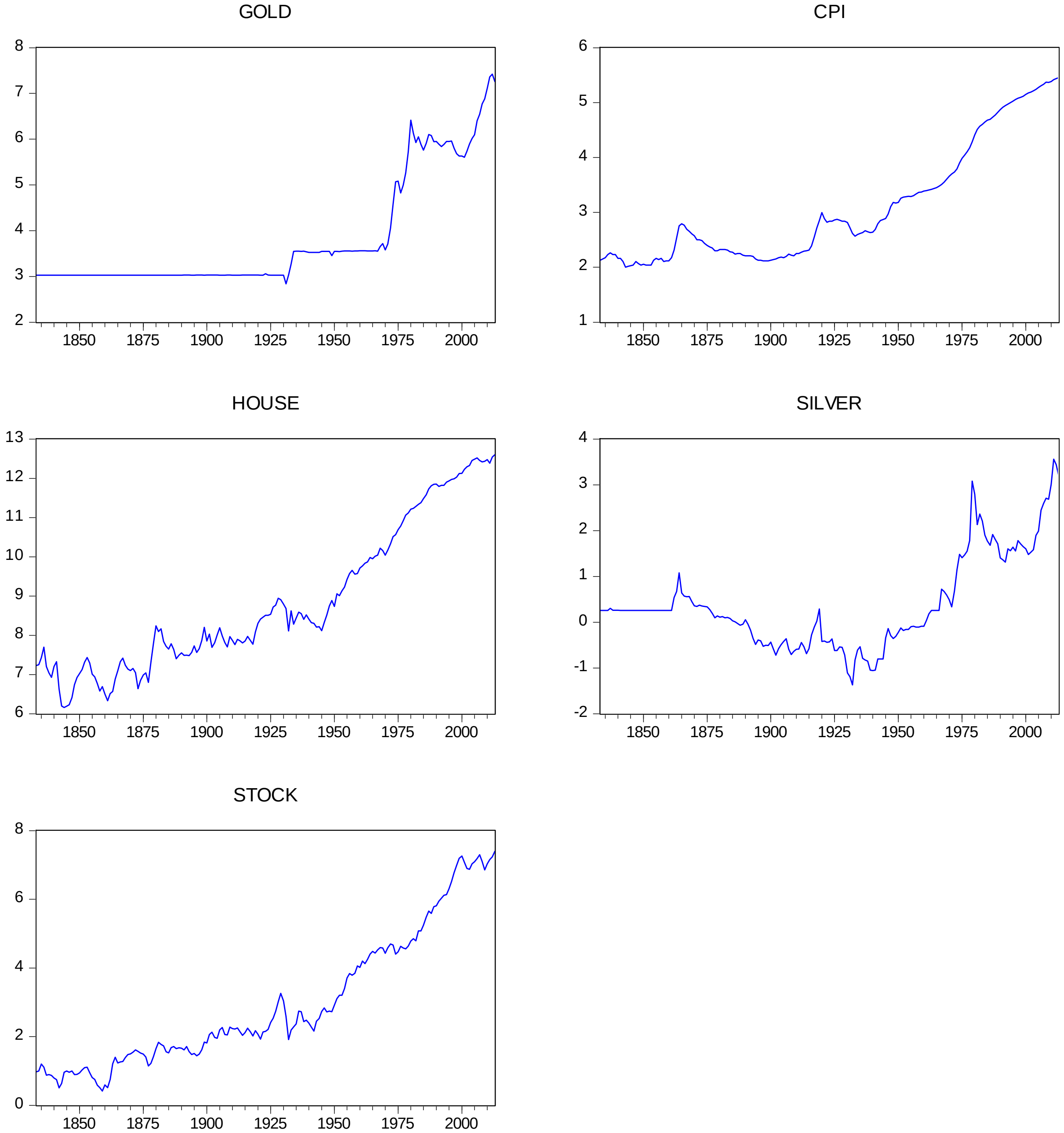

Figure A.1: Data plots in natural logarithms.