



On the economic determinants of the gold–inflation relation

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ARTICLE INFO

Article history:

Received 6 September 2013

Received in revised form

19 March 2014

Accepted 19 March 2014

Available online 4 May 2014

Keywords:

Gold

Inflation

JEL classification:

Q47

O13

C22

ABSTRACT

We examine the long term dynamic relation between inflation and the price of gold. We begin by showing that there is no cointegration between gold and inflation if the volatile period of the early 1980s is excluded from the data. However, we are also able to demonstrate that there is significant time variation in the relation, such that comovement between the variables has indeed increased in the last decade. Examination of the underlying macroeconomic factors that could generate time variation in the gold–inflation linkage suggests gold's sensitivity to inflation is related to interest rate changes, a finding that highlights the monetary nature of gold as a commodity.

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Introduction

Does holding gold protect an investor against inflation? Recent studies by Worthington and Pahlavani (2007), Wang et al. (2011) and Beckman and Czudaj (2013) all, in effect, say yes. Specifically, they all argue that the price of gold and the consumer price index (CPI) series are cointegrated. Gold and the CPI share a common long term trend. Such an outcome is therefore consistent with the view that gold should at least partially hedge inflation risk, lending support to the long held view that gold is a durable commodity that should underpin the world's monetary system. Nonetheless, the link between changes in the price of gold and inflation is not universal in the literature (see skeptical arguments in e.g. Blose, 2010; Baur, 2011; Erb and Harvey, 2013). In fact, Erb and Harvey (2013) argue that any supportive evidence on the linkage between gold and inflation stems from data in only one year, 1982.

Given the disparity in these findings, the main purpose of this study is to revisit the gold price–CPI relation. Our primary

contribution is that we document time variation in this relation and perhaps more importantly show that changes in gold's long term sensitivity to the CPI can be forecast by macroeconomic state variables, namely by interest rates. This is especially important to those market participants, who consider gold as an inflation hedge, and for those monetary authorities who use the price of gold as a leading indicator for inflationary pressures in the economy (see Ciner, 2011 and Shafiee and Topal, 2011).

Our investigation is motivated by both Atkeson and Ohanian (2001) and Stock and Watson (2006) who argue for a significant structural break in the U.S. inflation series in 1984, the onset of the so called “great moderation”. To avoid this issue, our data span the period between 1985 and 2012. We first show that there is no evidence of cointegration between the price of gold and the CPI, a finding consistent with Erb and Harvey (2013) and Blose (2010). This conclusion is supported by a battery of tests: the widely used system approach of Johansen (1991), the more recent, single equation error correction model based test advocated by both Pesaran et al. (2001) and Kanioura and Turner (2005), and the approach by Saikkonen and Lütkepohl (2000), which accounts for structural breaks in the variables.⁴

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⁴ The Saikkonen–Lütkepohl approach is also used by Worthington and Pahlavani (2007), who conclude that the cointegration between the price of gold and CPI is reversed when structural breaks in the variables are accounted for. This is in further detail below.

We then examine the temporal nature of the gold–CPI relation using a time varying regression in levels model based on the Kalman filter.⁵ The underlying intuition for this analysis is that while a cointegration framework primarily tests for a stable (stationary) relation between the variables there may be undetected (by cointegration) dynamics between the variables if there is time variation in the gold–CPI relation.

We find that the Kalman filter regressions do indeed point to time variation in the relation. Specifically, we show that the sensitivity of the gold price to CPI declines in the 1990s, which is why cointegration is rejected. However, the comovement between these two variables increases significantly again in the last decade. This naturally raises the question of explaining this time varying dynamic. For this purpose, we examine whether the change in covariation between gold and CPI can be explained by potential macroeconomic state variables. To the best of our knowledge, this is the first attempt to link the observed time variation to economic variables. Our findings are especially important for financial market participants because they provide insights in the inflation hedging ability of gold. We show that the United States (US) dollar has a significant and negative contemporaneous relation with gold's long term CPI beta. In other words, gold's sensitivity to inflationary risk tends to increase during dollar depreciations.

Finally, in addition to the contemporaneous linkages, we investigate whether changes in gold's long term CPI beta can be forecast by macroeconomic variables. In this sense we also add to a developing literature on the time dependence of these macroeconomic factors (e.g. Baur, 2011). Arguably, this could be important for financial market participants who consider gold as an inflation hedge. Granger causality tests for this part of the empirical analysis show that both short term (T-bill) and long term (T-bond) interest rates changes have significant predictive power for gold's CPI beta. In other words, gold's inflation sensitivity—hence its role as a hedge—can be forecast by interest rate changes. This further illustrates the monetary nature of gold as a commodity.

We organize the rest of the paper as follows: in the next section, we present the data set used in the study. In **Empirical findings**, we discuss the econometric analysis and the empirical findings. The final section includes the concluding remarks of the paper.

Data

The data set comprises monthly CPI and gold price series between January of 1985 and June of 2012, for a total of 319 observations. The CPI series are sourced from the Federal Reserve Bank of St. Louis Fred database and gold prices are the average monthly London PM fix sourced from the World Gold Council. Inflation is defined as the month on month percentage change. The series are plotted in Fig. 1, with gold plotted against the CPI level in the top panel and the rate of inflation in the lower. This Figure shows that gold prices were largely flat, or even declining, in the first part of our data, along with a noteworthy increase after approximately 2005. This is despite the CPI level showing a steady increase. Examining the inflation series we can see a volatile period at the start, with rising gold prices, a less volatile and flat gold period then commencing. This ends in early 1999, and we see here that with a short lag gold appears to respond, its slow decline stopping and a sharp increase in late 1999 marking the early stage of the great gold bull market. The deflationary period of early 2010

seems to have had little effect on gold prices. We choose to focus on the relation between the gold price and the US price inflation instead of a larger sample of countries primarily for two reasons. First, gold is priced in US dollars and the dollar remains the vehicle currency in international finance. Second, the correlations between the US and the G-7 country inflation rates are rather high, approximately .95 on average.

Following prior work, we take natural logarithms of the series in the empirical analysis to ensure better distributional properties. In Table 1, we report on the unit roots of the variables. Consistent with earlier studies investigating gold and financial prices (e.g. Ciner et al., 2013), both the Dickey and Fuller (1979), or Augmented Dickey Fuller (ADF) test and the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (1992) tests are supportive of the argument that the variables are nonstationary in levels and stationary in first differences. Hence, we can proceed with investigating the presence of cointegration in the series.

Empirical findings

Cointegration tests

We deploy three sets of cointegration tests. If this property is detected it can be argued that both gold and the CPI share a common stochastic trend. In turn this would imply that the two series will not arbitrarily drift apart in the long term. In other words, it can be argued that there is an inbuilt long term relationship. The first test is based on the system approach of Johansen (1991). In his approach, reduced form equations are estimated jointly; hence, exogeneity bias between the variables is minimized. This is the method largely used in prior work, such as Beckman and Czudaj (2013). Since the details of the Johansen approach to cointegration are discussed in many papers, for the sake of brevity we will omit them in the present study.

In the second approach, we also conduct the cointegration analysis using the more recent single equation error correction based approach based on papers by Pesaran et al. (2001) and Kanioura and Turner (2005). These authors, as well as Cook (2012), show that the conventional autoregressive distributed lag (ARDL) models can be transformed to test for the presence of long term relations in a single equation model, while reporting favorable power properties in their tests using simulation analyses. This approach offers the benefit of estimating both the long term and contemporaneous relations in a single equation model. However, it also requires exogeneity as a precondition.

Since the ARDL-based single equation testing method is relatively recent, it is briefly discussed below. Consider the following ARDL(1,1) model between gold and the CPI:

$$\text{Gold}_t = \delta_0 + \delta_1 \text{CPI}_t + \delta_2 \text{CPI}_{t-1} + \delta_3 \text{Gold}_{t-1} + \varepsilon_t \quad (1)$$

In this model, cointegration between Gold and CPI implies that $\delta_1 + \delta_2 \neq 0$ and $\delta_3 < 1$. However, since the variables contain unit roots, the model cannot be estimated using the standard least squares regression approach.

Both Pesaran et al. (2001) and Kanioura and Turner (2005) independently show that Eq. (1) can be expressed in an error correction format, which provides the added benefit of better economic interpretation while also facilitating hypothesis testing:

$$\Delta \text{Gold}_t = \alpha_0 + \alpha_1 \Delta \text{CPI}_t + \alpha_2 \text{Gold}_{t-1} + \alpha_3 \text{CPI}_{t-1} + \varepsilon_t \quad (2)$$

Testing for cointegration in this equation utilizes the notion that variables in levels should not appear in the model unless gold is cointegrated with the CPI. Hence, a rejection of $H_0 : \alpha_2 = \alpha_3 = 0$

⁵ An advantage of using the Kalman filter to examine this relation stems from the fact that it is robust to nonstationary data, which is the case for our gold price and CPI series as discussed further below. In other words, first differencing the variables, which would be required by the ordinary least squares (OLS), can be omitted and long run levels relations can be examined.

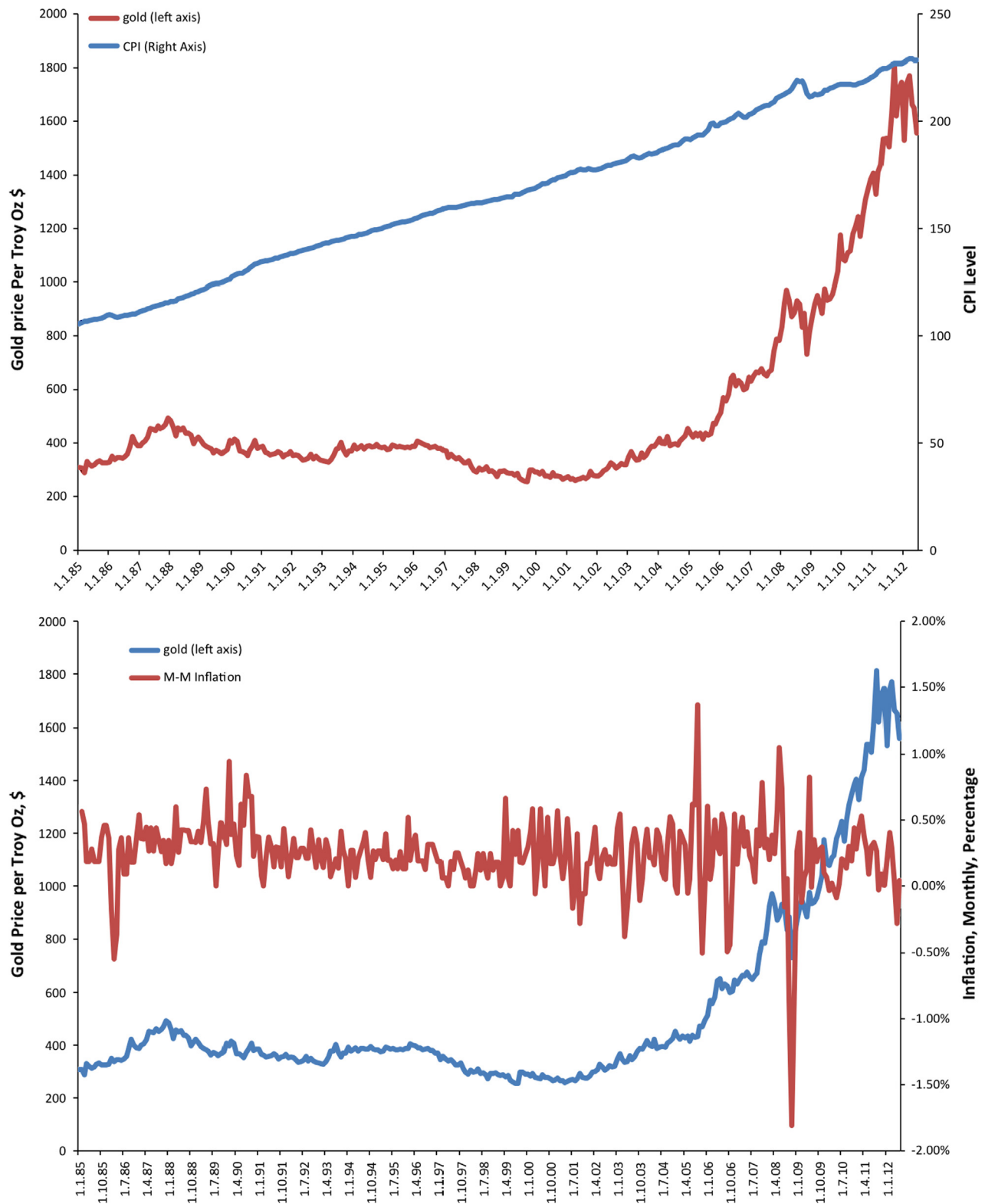


Fig. 1. Gold, inflation and CPI.

would indicate the presence of cointegration in the system. This can be tested by means of an F -test. Since the distribution is nonstandard, we use the response surface critical value generated by [Turner \(2006\)](#). As noted above, a potential issue with this approach is that it requires the exogeneity of the right hand side variable as a precondition. In an attempt to examine to what extent exogeneity issues cloud the analysis, we also conduct the

analysis using ΔCPI_t as the left hand side variable and calculate the cointegration tests in the same manner.

Finally, we test for cointegration using the method by [Saikkonen and Lütkepohl \(2000\)](#). The distinctive feature of their study is to present a framework to test for cointegration when there are possible structural shifts in the data, which could distort the findings of the conventional Johansen cointegration tests.

Table 1
Unit Root and Cointegration Tests.

Panel A: Unit Root Tests					
	Lag	ADF	KPSS		
CPI	2	−1.33	1.93		
Gold	2	0.31	2.35		
ΔCPI	1	−11.75	0.06		
ΔGold	1	−15.18	0.14		
Panel B: Johansen Cointegration Analysis					
Trace	CV	χ^2	Autocorrelation	ARCH	Non-normality
10.79	15.34	63.11 (.00)	13.17 (.86)	159.09 (.00)	535.33 (.00)
Panel C: Additional Cointegration Tests					
Single equation tests			Saikkonen–Lutkepohl test		
Exogenous Variable	F-test	CV	S–L Test	CV	
CPI	3.20	4.83	3.68	9.84	
Gold	3.22	4.83			

Note– This table provides the results of unit root as well as cointegration tests for the gold price and the CPI series. ADF refers to Augmented Dickey–Fuller tests; KPSS refers to Kwiatkowski–Phillips–Schmidt–Shin stationarity tests; and ARCH refers to Autoregressive Conditional Heteroskedasticity.

The Saikkonen–Lutkepohl approach is based on estimating the deterministic terms in the data by means of a generalized least squares (GLS) technique and subtracting the estimated deterministic terms from the observations. A dummy variable is also included

to address structural shifts in the mean. Then, the Johansen type, reduced rank regression methodology, can be applied to the adjusted series, with critical values for the test statistics reported. As mentioned above, this is the method used by [Worthington and Pahlavani \(2007\)](#), who also provide a detailed exposition of this method, to test for cointegration between gold prices and the CPI.

We first report the Johansen cointegration test statistics, in Panel B of [Table 1](#). The lag lengths used in the models are calculated by the Akaike Information Criterion (AIC). In the Johansen analysis, an important issue is to determine how to treat the trend in the system as the critical values differ depending on this specification. We examine, by means of the chi-square test described in [Johansen \(1991\)](#), whether the trend in the VECM could be restricted to enter the system only via the error-correction variable or the alternative hypothesis of the trend being orthogonal to the cointegration vector.

The chi-square test reported in Panel B of [Table 1](#) indicates that this restriction can be comfortably rejected. This implies that the trend should be estimated orthogonal in the system ([Johansen's, 1991](#) Case 3). We also test for the specification of the system using multivariate tests for autocorrelation and ARCH effects in the VECM residuals. It is observed that the null hypothesis of no autocorrelation cannot be rejected. However, there is evidence for ARCH effects in residuals and hence the normality assumption is also rejected. Since the existing literature suggests that the Johansen procedure is robust to ARCH effects in residuals ([Lee and Tse, 1996](#)), we proceed with the estimated VECM. The resulting trace test indicates no evidence for cointegration between the series.

Our second approach to testing for cointegration between the price of gold and the CPI is the single equation ARDL based cointegration tests. These are reported in Panel C of [Table 1](#). Again, the conclusion is the same; the null hypothesis of no cointegration cannot be rejected. Furthermore, we find that the results seem robust to potential exogeneity issues in the single equation testing framework as the conclusions are not affected by which variable is used in the left hand side of the equation.

The Saikkonen–Lutkepohl tests for cointegration are presented next. As discussed above, this approach is built on estimating GLS

regressions of the data series to adjust for deterministic terms, including possibly structural shifts, and then, testing for cointegration in a Johansen type framework. Thus, we first examine whether any structural shift can be detected in our series. We use the approach suggested by [Lanne et al. \(2002\)](#) and indeed find that both series have a structural break in November of 2008.⁶ At this junction, it might be useful to compare our findings with those of [Worthington and Pahlavani \(2007\)](#).

These authors also use the Saikkonen–Lutkepohl method but their data set spans the period between 1945 and 2006. They detect structural breaks in gold series in January of 1973 and December of 1979 and for the CPI, in February of 1973 and January of 1979. Since our sample covers a later time period, we can avoid those breaks. However, we also have approximately six more years of data than their study, in which we actually find the single structural break in November of 2008. Hence, our results in this part are consistent with their analysis. However, the actual cointegration results point to a different direction as our Saikkonen–Lutkepohl test statistic, reported in Panel C of [Table 1](#), is insignificant.⁷ In sum, all three cointegration approaches used in the analysis consistently reject a stable long term relation between gold and the CPI.

Furthermore, we examine the evolution of the cointegration property between the gold price and the CPI. As reported in the introduction, several recent papers actually report evidence of cointegration between these variables, in contrast to our analysis. The main difference between the aforementioned papers that find cointegration and our analysis is that we do not include data prior to 1985. To examine the evolution of cointegration, we recursively estimate the VECM and plot the largest eigenvalue that is used to calculate the Johansen trace test for cointegration. We plot our findings in [Fig. 2](#). It can be observed from this figure that there is a noteworthy drop in the eigenvalue after the mid-1980s. This finding is consistent with our intuition that the reported cointegration between these variables is driven by data from the late 1970s and early 1980s.

⁶ It is noteworthy that November of 2008 is the beginning of the US Federal Reserve's unconventional monetary policy. Specifically, it was announced on November 25, 2008 that the Fed would start purchasing mortgage backed securities (MBS).

⁷ These test statistics are estimated using the JMulti software.

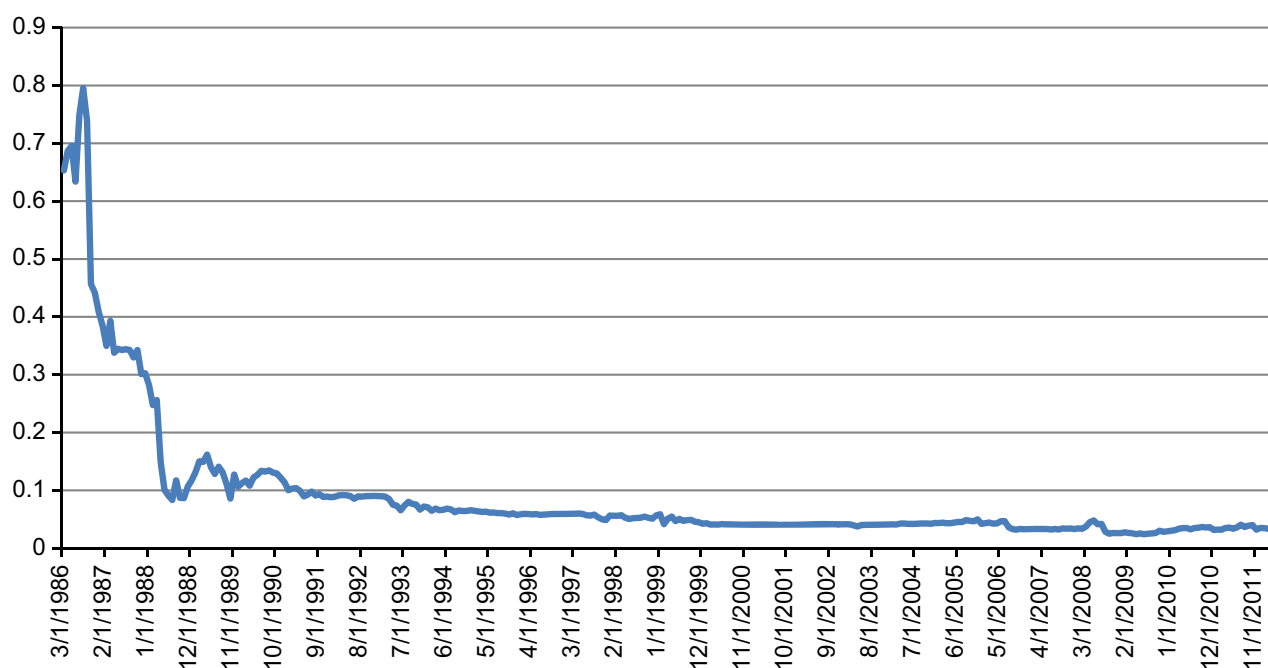


Fig. 2. Time variation in eigenvalue.

Time-varying gold–CPI dynamics

We proceed by further examining the time variation in the sensitivity of gold prices to the CPI. While the cointegration analysis can very usefully detect the long term equilibrium relation between the variables, it is inherently designed to capture stable linkages. If two series exhibit divergence at the beginning of the sample, but have started moving together in the later part of the sample, cointegration will likely be rejected. Thus, the cointegration methodology may not be suitable when there are dynamic structural changes in the series under investigation. Other authors, including Aroui et al. (2012) have shown that structural breaks have definitely occurred with respect to the gold price. Therefore, while the results in the previous section show that cointegration can be comfortably rejected in our dataset, there might be uncovered changes in the long term relation between the price of gold and inflation. Investigation of this phenomenon forms the focus of the empirical analysis in this part of the study.

To accomplish this goal, we consider the following model:

$$\text{Gold}_t = \beta_0 + \beta_{1,t} \text{CPI}_t + u_t, \quad u_t \sim N(0, \sigma_u^2) \quad (3)$$

$$\beta_{1,t} = \beta_{1,t-1} + \theta_t, \quad \theta_t \sim N(0, \sigma_\theta^2) \quad (4)$$

In this approach, the sensitivity of the gold price to the CPI is captured by $\beta_{1,t}$, which is modeled as a random walk process and where a smooth structural change in the linkage between variables is permitted. The model also fits naturally into the state space framework with the state being represented by the time varying coefficient. Note that the variables are used in levels form in the equation because our goal is to examine long term dynamic linkages.⁸

As shown in the previous section, both the CPI and gold price series are nonstationary in levels. However, an important advantage of using the Kalman filter is that it is robust to nonstationarity in the data. This is due to the distributions of the state variables, which are dependent upon the previous observations. In other words, the usual stationarity assumption of the ordinary least squares (OLS) is not required since all parts of the model are permitted to vary and the model is reestimated recursively. This approach is discussed in greater detail in Bornhöff (1992) among others. The diffuse method of de Jong and Chu-Chun-Lin (2003) is used to initialize the filter.

The plot of gold's CPI beta ($\beta_{1,t}$) may be found in Fig. 3, which supports the view that there is time variation in the gold–CPI relation as hypothesized a priori. Specifically, the 1980s and 1990s are characterized by a decline in the sensitivity of the gold price to inflationary shocks, which is consistent with the lack of support for cointegration between those variables detected in the previous section.

However, in the latter part of the sample, specifically after 2002, there is an increase in the long term comovement between the two variables. The findings of this analysis illustrate the usefulness of examining smooth structural change in the relation. Increased sensitivity to inflation indicates that data in the most recent decade would imply gold to be a hedge against inflation, while this conclusion cannot be reached in data from the earlier era.

In the beginning of the sample the CPI beta of gold was as high as 2.5, which suggests that the gold price was very sensitive to changes in inflation. However, throughout the 1990s, the beta falls and in fact, at the end of the 1990s, the CPI beta for gold price is estimated as almost zero. Later, the gold beta starts an upward trend from the 2000s again almost reaching the level of 2.5 at the end of the sample period in 2012. From highly sensitive in the 80s through a period of low sensitivity in the 1990s gold has now returned to being highly sensitive to inflation.

⁸ It is common in financial research to conduct the empirical analysis using returns of assets, which are obtained by first-differencing the series. To a large extent, this is due to the fact that ordinary least squares regression requires stationarity in data and financial time series are nonstationary (see Baxter (1994)

(footnote continued)

for a discussion). The Kalman filter is robust to nonstationarity of the data and therefore, has important advantages over the OLS.

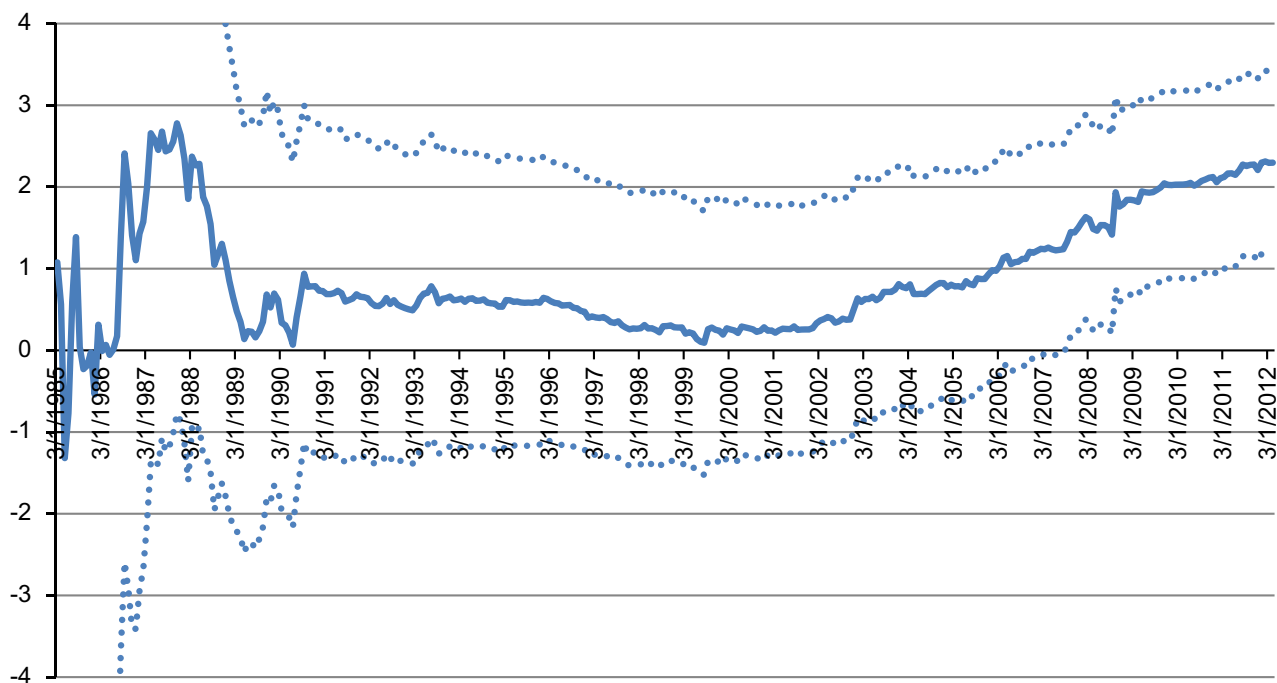


Fig. 3. Time variation in gold–CPI –elation ($\beta_{1,t}$).

Economic determinants

An important question following the above empirical analysis is what causes the observed variation in the gold–CPI relation. An advantage of using the Kalman filter to model the parameter of interest is that it generates a continuous series of estimates and therefore, macroeconomic state variables can be considered and statistically tested for explanatory power in the relation. If we detect variables that can forecast the CPI sensitivity of gold, this would presumably be of significant interest for market participants, who need to adjust their holdings in anticipation of inflationary shocks.

Consistent with Baur (2011) we consider a number of explanatory variables for analysis: the US 3-month Treasury bill (T bill) rate, the 10-year Treasury bond (T-bond) rate, the trade weighted (TWI) US dollar index, industrial production and the University of Michigan consumer confidence indexes. We include interest rates to capture changes in the yield in the economy. The industrial production index captures changes in the business cycle, while the consumer confidence index accounts provides a forward looking estimate of sentiment and is much used by financial market participants. The US dollar index accounts for the fact that several studies mentioned earlier argue that gold should primarily be considered a hedge against the US dollar.

We first examine the contemporaneous relation between gold's CPI beta and our macroeconomic variables by estimating the following regression:

$$\Delta\beta_{1,t} = \pi_0 + \pi_1\Delta X_t + \varepsilon_t \quad (5)$$

In this equation Δ represents the first difference operator and X_t represents the macroeconomic state variables discussed above. The results of this regression are presented in Table 2.

Consistent with Baur (2011), the findings show that only the US dollar index has significant explanatory power for time variation in $\beta_{1,t}$ and importantly has a negative sign. This finding is consistent with studies that suggest that gold should be considered as an alternative to paper currency (e.g. Baur, 2011; Ciner et al., 2013

Table 2

Contemporaneous Linkages.

	USD	Ind. prod.	Sentiment	T-bill	T-bond
π_1	−2.84 (0.02)	−1.27 (0.37)	−0.23 (0.21)	−0.03 (0.71)	0.11 (0.18)

Note: This table provides the estimation results for Eq. (5), which tests for contemporaneous relations between changes in gold's CPI beta and various macroeconomic state variables: US 3-month Treasury bill (T bill) rate, the 10-year Treasury bond (T-bond) rate, the trade weighted US dollar index (USD), industrial production (Ind. prod.) and the University of Michigan consumer confidence indexes (sentiment). The bivariate regression coefficients and their p -values (in parenthesis) are reported. P -values are calculated using White's robust standard errors.

among others). It also seems plausible that dollar depreciations are associated with higher price inflation in the US, which would increase the demand for gold as a hedge. This scenario would create increased comovement between the price of gold and inflation, which is of course detected in our data.

Finally, we examine whether gold's CPI beta can be forecast. As mentioned, this is arguably important for financial market participants who could use gold as a potential inflation hedge. We use the same set of macroeconomic variables discussed above and conduct Granger causality tests within the context of bivariate vector autoregression (VAR) models. Each variable is treated as endogenous in a VAR model and causality tests are based on regressions of each variable on autoregressive lags of its own as well as other variables in the system. Since many references are available on the use of VARs in financial research, we do not discuss them in detail in this study. An important issue in the VAR analysis, however, is determination of the lag length to be used. We rely on the AIC in this analysis, which tends to produce longer lags than Schwarz Bayesian criterion (SBC).

The causality test results are reported in Panel A of Table 3. The findings indicate that both short- and long-term interest rates have statistically significant predictive power for gold's CPI beta at statistically comfortable levels (p -values $< .05$). This further

Table 3
Forecasting Gold's CPI Beta.

Panel A: Causality Tests		F-test	p-Value
	Lag		
USD	7	1.79	(.08)
Ind. prod.	7	.94	(.46)
Sentiment.	7	.58	(.76)
T-bill	6	2.47	(.02)
T-bond	10	3.46	(.00)
Panel B: Impulse Response Functions		T-bond	
	T-bill		
1	-.034 (-.030, -.011)	-.000 (-.016,.026)	
2	-.036 (-.067, -.008)	-.019 (-.058,.013)	
3	-.021 (-.053,.009)	-.016 (-.051,.017)	
4	-.023 (-.058,.011)	-.008 (-.047,.026)	
5	-.042 (-.080, -.038)	-.024 (-.065,.013)	
6	-.054 (-.094, -.016)	-.048 (-.089, -.005)	
12	-.073 (-.123, -.021)	-.047 (-.090, -.000)	
24	-.082 (-.140, -.023)	-.050 (-.095, -.002)	

illustrates the role of gold as a monetary asset. The US dollar exchange rate on the other hand can be considered a statistically significant predictor only at 8% significance level.

We also report the accumulated impulse response functions along with their critical values for both interest rates on gold's inflation beta in the bottom Panel B of Table 3. The critical values are calculated by Efron's (1979) bootstrap method with 1000 replications. As expected the impulse response functions show that a negative relation exists between interest rate changes and gold's CPI sensitivity. In other words, declines in interest rates increase the importance of inflation for the gold price. This is consistent with the notion that a monetary easing stokes fears of higher inflation, which in turn impacts the price of gold.

Conclusions

This article provides further evidence on how the cointegration relationship between the gold price and the CPI disappears in data after 1985, contrary to the evidence presented in several recent papers. Overall, we show that cointegration analysis may be less useful if there are significant time varying dynamics in the system.

To account for this possibility, we employ a Kalman filter based approach to investigate the dynamic relation between the price of gold and the CPI. Our analysis shows that a stable link between these variables does not exist. In particular, we find that gold's CPI beta becomes gradually smaller throughout the 1990s, which is the underlying reason for rejection of cointegration. On the other hand, in the 2000s, we observe a significant increase in gold's beta, which is perhaps why there are conflicting empirical results on the relation between the price of gold and inflation in the literature.

As it is important for financial market participants to determine when gold could play the role of an inflation hedge, we test whether time variation in gold's CPI beta can be explained by various macroeconomic state variables. We find that a contemporaneous, statistically significant, negative relation exists between changes in gold's inflation beta and the value of the US dollar. This finding can be regarded as evidence for the notion that gold represents an alternative to the paper currency with its historical role at the center of the monetary system.

Perhaps of greater importance for market participant is whether gold's CPI beta can be forecast. This would help to determine when gold could play its role as an inflation hedge. For this purpose, we

conduct Granger causality tests between gold's CPI beta and our macroeconomic state variables and report evidence to suggest that both short and long term interest rate changes forecast changes in gold's beta. The impulse response function analysis of the VAR models further show that the impact is negative and the cumulative impact of interest rate changes are felt in approximately 6-months. This finding, of course, further illustrates the monetary nature of gold as a precious commodity. Therefore, one key insight from this research is that we are able to provide an answer to the question of when gold may act as an inflation hedge.

Our results also have implications for policy makers, chief among them central bankers of emerging market economies. During the global financial crisis of 2007–2010, these institutions actually increased their holdings of gold as reserve assets, presumably under the assumption that gold would preserve its purchasing power, while also providing a hedge against inflationary shocks. The empirical analysis undertaken in this article cautions against such a policy. It is also noteworthy that central banks of developed economies have in fact been reducing the share of gold in reserve assets. These actions are more consistent with the findings of the present article. Moreover, a direct implication of our analysis is that changes in the price of gold should not necessarily be interpreted as signs of changes in the inflationary pressures present in the economy. There is a school of thought in monetary policy making that suggests commodity prices, and primarily gold, can be used to set monetary policy objectives. Our analysis also cautions against this viewpoint.

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