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Are common stocks a hedge against inflation in emerging markets?

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Abstract This paper examines the inflation-hedging ability of common stocks in the long run for emerging market countries using monthly data on stock and goods prices over the period 1982:01–20,016:01. Johansen's (J Econ Dyn Control 12:231–254, 1988) method of cointegration is employed for 28 countries for which the order of integration of the underlying series is the same, and Pesaran et al.'s (J Appl Econom 16:289–326, 2001) autoregressive distributed lag (ARDL) bounds test is used for 18 countries for which the order of integration is not the same. In only 10 cases is the long-run relationship established between stock and goods prices when the former test is used, whereas such a relationship is established in 7 cases when the latter test is used. The results of error-correction representations normalized on stock prices indicate that stock prices take a long time to return to their long-run equilibrium relation with goods prices. Overall, common stocks provide a good hedge against inflation in the long run in more than one-third of the cases examined. One implication that emerges from these results is that in the majority of the countries, monetary authorities are not able to control inflation by reducing the nominal interest rate. Another implication is that the monetary growth is not the key factor determining the long-run inflation rate in these countries.

Keywords Fisher hypothesis · Inflation hedge · Stock prices · Cointegration and emerging markets

JEL classification G10 · G15 · C32 · E44 · E31

1 Introduction

Since the Great Inflation, which lasted from 1965 to the mid-1980s, academics and practitioners have been quite concerned about the inflation-hedging property of

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common stocks. That common stocks should have the ability to hedge against inflation is a proposition that is typically rationalized on the grounds that common stocks represent residual ownership of physical production facilities whose values increase with rising prices (inflation). Over the last five decades, a large body of theoretical and empirical work has accumulated in the literature to test the validity of the inflation hedging capability of common stocks.[1] A common feature of this work is that it explores the relationship between nominal returns on the underlying assets and inflation in the context of the Fisher hypothesis such that if the hypothesis holds, then the prices of these assets should move in a one-to-one positive relation with goods prices and, as a consequence, the expected nominal returns on them will be equal to expected inflation rates over time. However, results emerging from this work have generally been inconclusive, giving rise to controversy among researchers about both the validity of the underlying hypothesis and the inflation-hedging ability of common stocks.

One implication of the invalidity of the Fisher relation is that real returns on financial assets such as common stocks are unlikely to be independent of inflation rates, eventually resulting in non-neutrality of monetary policy. In reality, this hypothesis is one of the building blocks of the new classical macroeconomic model stating that over the long run, inflation is driven exclusively by the monetary growth, and as such, real return and output levels tend to return to their natural levels determined by real (non-monetary) forces. The second implication is that monetary authorities are unable to control inflation by reducing nominal interest rates. The third implication is that the inability of financial assets to hedge against inflation creates uncertainty across financial markets, thereby adversely affecting investment and saving decisions in an economy. The final implication is that investors around the world are unable to improve the inflation protection properties of their equity investment portfolios through global portfolios, since inflation seldom affects all regions of the world at the same time.

The objective of this paper is to examine the inflation-hedging ability of common stocks for emerging market countries by examining the proposition embedded in what is termed "the generalized Fisher hypothesis" (GFH) in that stock prices should move positively in a one-to-one relation with goods prices and, hence, expected nominal returns on stocks will be equal to inflation rates over time. The motivation for reinvestigating the inflation-hedging ability of common stocks for emerging market countries goes as follows. First, historically high average expected returns and volatility of equity stocks have attracted the attention of many researchers (among others, Barnes et al. 1999; Choudhry 2001; Al-Khazali and Pyun 2004; Alagidede 2009; Alagidede and Panagiotidis 2010) to testing how the GFH performs for the stock markets of emerging market countries and whether emerging equity markets protect their investors against inflation. Second, new equity markets of the emerging market countries have provided the world's investors with a larger and an increasingly important set of investment opportunities for global portfolio diversification (Harvey 1995).[2] Third, in fact, inflation risk has been relatively more intense

[1] An enormous work has also been carried out on the inflation hedging ability of other financial (e.g. bonds and Treasury bills) and real assets (e.g. gold and real estate). For a recent, detailed and comprehensive survey on this issue, see Arnold and Auer (2015).

[2] It is argued that average expected returns and their volatility have historically been higher in equity markets of the emerging market countries. However, given the low correlation between equity returns of developed and emerging market countries, the inclusion of emerging market assets in a mean-variance efficient portfolio is likely to significantly reduce portfolio volatility and increase expected return.

in emerging market countries than in developed countries, since inflation has always been much higher in these countries due to higher interest rates and food prices on the one hand and the much greater importance of food and stronger economic growth rates on the other hand (Amenc et al. 2009). High sustained rates of inflation do not only make investors more risk averse by eroding the purchasing power of nominal returns they obtain over the holding period but also exert a detrimental effect on an economy's long-run real economic activity by adversely affecting capital formation in the long run (Barnes et al. 1999). Fourth, episodes of high inflation rates lasting over many quarters often cause crises (including currency crisis) that adversely affect the real economy. Therefore, it is important for investors around the world to know how stock markets perform in the emerging market countries and whether they provide them hedges against inflation risk.

This paper intends to make the following contributions to the literature on the GFH and the inflation-hedging ability of commons stocks for emerging market countries. First, the paper utilizes an extended dataset for 50 countries, including countries in Europe, Asia, Middle East, Latin America and Africa, some of which have not been examined yet to the best of our knowledge. Second, we use a larger sample that includes both the pre- and post-2008-2010 periods of the international financial crisis and recession for most of the emerging equity markets.[3] It is worth noting that the sample includes some countries that had undergone additional (local or regional) crises over the period under consideration (e.g., the Asian currency crisis of July 1997, Brazilian real crisis of January 1999, Mexican peso crisis of December 1994 and Russian default crisis of August 1999). Third, although regression analysis is also performed to examine whether nominal returns are positively related to inflation rates in emerging market countries, we employ cointegration analysis to investigate the long-run relationship between stocks and goods prices to determine whether stock markets in emerging market countries provide any hedge against inflation. The remainder of this paper is structured as follows. Section 2 provides a brief account of a review of some selected studies focusing on the empirical validity of the GFH for developed market and emerging market countries. While section 3 is devoted to discussion on the data employed and the model and methodology used, section 4 presents and interprets the results. Section 5 concludes the findings and derives policy implications.

2 Empirical evidence

2.1 Evidence for developed market countries

Motivated primarily by the surge of inflation in the 1970s, earlier empirical work explores the stock-return and inflation relation for highly developed stock markets of major industrialialized (or OECD) countries and whether common stocks in these markets have the ability to hedge against inflation. The work in this context has generally followed four main strands. The first strand relates to the earliest empirical

[3] The exceptions include such countries as Bahrain (2007:08–2016:01), Qatar (2009:01–2016:01), United Arab Emirates (2007:01–2016:01), Slovenia (2006:04–2016:01), Chile (2009:01–2016:01) and Venezuela (2007:12–2016:01) for which the sample starts from the period when the underlying crises had already erupted but not ended.

studies carried out by, inter alia, Oudet (1973), Lintner (1975), Nelson (1976) and Fama and Schwert (1977) on the validity of the GFH for the U.S. stock market. Notwithstanding its wide acceptance, these studies have reported evidence that is generally inconsistent with this hypothesis, showing that nominal common stock returns are negatively related to both expected and unexpected inflation rates during the post-1953 period. This evidence, which has been interpreted as "the stock return-inflation puzzle" in the literature (see, for example, Gallagher and Taylor 2002), is not supported for the U.K. by Firth (1979), who shows that British stocks, unlike U.S. stocks, provide some hedge against inflation. Similar results are reported by Gultekin (1983a) using data from the Livingston survey of expectations and showing that the GFH holds much better for *ex ante* expectations than for *ex post* realization.

The second strand involves studies extending the empirical testing of the GFH to industrialized countries other than the U.S. and U.K. to evaluate whether or not the stock return-inflation puzzle is applicable to other major industrialized countries. Cohn and Lessard (1981) argue that in reality, the negative relationship between stock returns and inflation is characteristic of most major industrialized countries. Testing the GFH for 8 major industrialized countries – Canada, France, Germany, Italy, Japan, the Netherlands, the U.K. and the U.S. – using quarterly data over the period 1971:01–1979:04, they obtain results that are generally consistent with those of the earliest studies, namely, that common stocks are poor hedges of inflation. Similar results have been reported by Solnik (1983) and Cohn and Lessard (1981) who present findings that suggest that the inability of stock prices to keep up with the general price level is not a phenomenon restricted to the U.S. alone. Solnik (1983) tests the GFH using monthly data for 9 OECD countries – Belgium, Canada, France, Germany, Japan, the Netherlands, Switzerland, the U.K. and the U.S. – over the period 1971:01–1980:12 and reports results that confirm the negative relationship between nominal stock returns and inflation rates in all countries, with the exception of Canada. Gultekin (1983b) tests the GFH for 26 OECD countries, including the U.S., and finds that in most countries, there is no consistent positive relationship between stock returns and inflation.

The third strand relates to some studies (Barnes et al. 1999; Ahmed and Cardinale 2005; Li et al. 2010) suggesting that the inflation-stock returns nexus varies across different inflationary regimes. Barnes et al. (1999) explore whether common stocks' ability to provide a hedge against inflation varies across different inflationary regimes. They examine the GFH for 25 countries with low and high inflation rates and find results indicating that the underlying relation performs better in the latter regime than in the former regime, and hence, common stocks provide a better inflation hedge in high inflation countries. They obtain results that are inconsistent with the GFH for 16 countries with low-to-moderate inflation rates and in which the inflation coefficients are not only negative but are also statistically insignificant in all cases, except the U.S. Only in 9 countries do they find that the inflation coefficients are positive. Of these 9, in 6 countries (Austria, Germany, India, New Zealand, the Philippines and the U.K.) the inflation coefficients are insignificantly positive, whereas in 4 cases (Chile, Israel, Mexico and Peru) the inflation coefficients are significantly positive. Ahmed and Cardinale (2005) explore the relationship between inflation rate and stock returns in an inflationary regime framework and show that on the U.K. market, over the period 1900–2002, the inflation rate did matter for equity returns. In contrast, Li et al. (2010) examine whether the cross-sectional relationship between inflation and stock returns varies in an inflationary regime framework. They investigate the relationship between inflation and stock returns in the short term and medium term and under different inflationary regimes. Using monthly data on the retail price index, the FTSE all-share index, and 10 industry indexes, namely, oil and gas, basic materials, consumer goods, health care, consumer services, telecoms, financials, information technology and utilities for the U.K., they obtain results indicating that although U.K. stocks fail to hedge against inflation in the short run, evidence is mixed in the medium term. Moreover, results from different inflationary regimes show that the relationship between inflation and stock returns varies across different regimes.

The fourth strand relates to research work that centers around the rehabilitation of the GFH by looking into some neglected aspects of the empirical testing. The empirical testing focusing on the rehabilitation of the GFH explores the following issues: (i) long-run data consisting of a century or two and a long holding period, (ii) more sophisticated estimation procedures, including cointegration analysis, panel cointegration, non-linear cointegration, and quantile regression analysis and (iii) the index of the highest beta stocks. Boudoukh and Richardson (1993) argue that the studies reporting the failure of common stocks as inflation hedges often use data on stock returns and inflation rates over short periods and short investment horizons. Since the GFH hypothesis is an equilibrium condition that is expected to hold over a long period, it is important to examine how stock returns move with inflation rates over longer investment horizons and longer periods for two reasons. First, from a practical perspective, many investors hold stocks over longer horizons. Second, the relationship between stock returns and inflation rates over longer horizons is of particular importance given that the results at short horizons appear to be inconsistent with the underlying hypothesis. Using an approximately two-century annual dataset on inflation and stock returns over the period 1802–1990, Boudoukh and Richardson (1993) find that over a long horizon, nominal stock returns are positively related to both *ex ante* and *ex post* inflation rates. They note that these results are somewhat robust with respect to particular sub-periods over the past two centuries as well as to the choice of U.S. or U.K. markets. Using stock returns over 1- and 5-year holding periods and different long sample periods (1802–1890, 1870–1890 and 1914–1990), Boudoukh and Richardson (1993) report that the relationship between stock returns and inflation is positive when a 5-year holding period is used rather than a 1-year holding period. The 5-year relations yield positive and significant coefficients ranging between 0.38 and 2.12, with most coefficients exceeding 1. Similar results are reported for the U.S. market in an earlier study conducted by Miller et al. (1976), who investigate an extended period using nearly a century of data from 1876 to 1971. Miller et al. (1976) note that annual stock returns are negatively correlated with annual inflation rates over a short run, whereas the underlying relation is positive over the much longer period of 1875–1970. They also estimate a long-run elasticity of 0.50 (which is not significantly different from one) using annual stock return and inflation rate data over the period 1875–1970. Solnik and Solnik (1997), Lothian (1998), Schotman and Schweitzer (2000) and Lothian and McCarthy (2001) also find that investment in common stocks provides a hedge against inflation only in the long run. Solnik and Solnik (1997) test the GFH for 8 countries and show that the inflation coefficient approaches unity as the investment horizon increases. In contrast, Lothian (1998) and Schotman and Schweitzer (2000) use annual data for a panel of 23 OECD countries over the period 1973–1994 and find that investment in common stocks provides a hedge against inflation only over the long run, that is, when the investment horizon exceeds 15 years. Lothian and McCarthy (2001) reexamine the GFH using data on stock and goods prices for 14 OECD countries over the post-World War II period and compare the behavior of the underlying series in that sample with those of the U.K. and the U.S. over the long period, 1790–2000. Contrary to many studies over the past several decades, they find that movements in stock prices keep pace with movements in goods prices only over longer periods. The authors conclude that the reason why common stocks failed to hedge inflation in previous studies is that the relationship takes quite a long time to appear. The results are, however, in sharp contrast with findings by Alagidede and Panagiotidis (2012), who apply a quantile regression framework[4] to test the relationship between contemporaneous stock returns and inflation rates in the G-7 countries – Canada, France, Germany, Italy, Japan, the U.K. and the U.S. – over the period 1970:01–2008:04. The results indicate that there is a positive relationship between stock returns and inflation rates in almost all cases.

One problem with the studies attempting to rehabilitate the GFH over the long run is that they invariably test the hypothesis using a specification that essentially represents a short-run, not a long-run, relation between stock and goods prices. This is because they employ first-difference data on the underlying variables regressing nominal stock returns on inflation rates, which is essentially a short-run relationship. A number of researchers, inter alia, Ely and Robinson (1997), Anari and Kolari (2001), Al-Khazali and Pyun (2004), Alagidede and Panagiotidis (2010) and Hassan et al. (2015) use models based on level data on stock and goods prices and cointegration analysis that allow them to test whether stocks maintain their value relative to goods prices in the long run. Ely and Robinson (1997) argue that a large body of evidence indicating that stock markets tend to perform poorly during inflationary time periods is mostly derived from models structured to estimate the short-run relationships between stock returns and inflation rates. They use a reduced form approach and cointegration analysis to explore the relationship between stock and goods prices for 16 industrialized countries using quarterly data over the period 1957:02–1992:03 and test whether stock prices maintain their value relative to goods prices and whether these response patterns depend on the source of inflation shocks. They use vector error-correction (VEC) models that allow them to test the long-run relationship between stock and goods prices by incorporating real output and money variables that have been suggested as playing a role in determining the relationship between stock return and inflation and to estimate the impulse response functions to assess the response of stock and goods prices to innovations in both the money supply and real output. Employing Johansen's (1988) multivariate vector autoregressive (VAR) cointegration technique, they obtain evidence rejecting the null hypothesis of no cointegration between stock prices, goods prices, real output and money supply only in a minority of cases. Even in many of these cases in which the authors find cointegration, there is no evidence for a one-to-one correspondence between stock and goods prices. However, based on what they term as their 'principle results' of VEC models, they conclude that stocks do maintain their value relative to movements in overall price indexes (CPI, PPI and GDP deflator) and

[4] The quantile regression framework, which was developed by Koenker and Bassett (1978), provides estimates of the linear relationship between the regressors and a specified quantile of the dependent variable. For a detailed analysis of quantile regression, see Koenker and Hallock (2001).

that this conclusion generally does not depend on whether the source of inflation shock is real or monetary in nature. However, these results are in sharp contrast with those obtained by Anari and Kolari (2001),[5] who use a similar methodology and monthly data on stock and goods prices for 6 industrialized countries (Canada, France, Germany, Japan, the U.S. and the U.K.) over the period 1953–1998. Anari and Kolari (2001) produce results showing that there is a long-run relationship between stock and goods prices and that the estimates of the long-run elasticities of stock prices with respect to goods prices generally exceed 1, ranging between 1.04 (France) and 1.65 (Japan), and supporting the GFH in all cases. The authors argue that the coefficients of the GFH (ranging from 1.04 to 1.65) are more consistent with those reported by Boudoukh and Richardson (1993) than those reported by Miller et al. (1976). They also estimate an error-correction model and find that the estimated coefficients of speed of adjustment lie between 0.01 and 0.03, implying that it takes a very long time for stock prices to return to their long-run relation following unexpected movements in goods prices. Similar results are obtained by Al-Khazali and Pyun (2004), Luintel and Paudyal (2006)[6] and Alagidede and Panagiotidis (2010), who find evidence of a positive long-run relation between stock and goods prices for 9 Pacific Basin countries, the U.K. and 6 African countries, respectively. Kim and Ryoo (2011) also find that U.S. common stocks have been a hedge against inflation since the early 1950s. They test the long-run relationship between stock and goods prices using a century-long U.S. monthly dataset over the period 1900:01–2009:06 and different window lengths (10, 20, 30 and 40 years). Using the Seo (2006) test for cointegration in a two-regime threshold vector error-correction model (TVECM),[7] they show that stock and goods prices are cointegrated with unit elasticity, with stock returns and inflation showing asymmetric error correction.

However, results obtained by Hassan et al. (2015) are somewhat unsupportive of the GFH. They apply alternative cointegration procedures, such as the maximum likelihood technique of cointegration of Johansen and Juselius (1990), the autoregressive distributed lag (ARDL) method of cointegration of Pesaran et al. (2001) and the non-linear two-regime threshold cointegration procedure of Hansen and Seo (2002), to data for 19 OECD countries. The results based on the threshold cointegration technique confirm cointegration in five countries (Ireland, Italy, Japan, the Netherlands and Switzerland), whereas the linear cointegration method of Johansen and Juselius (1990) produces cointegration in only three countries (Finland, Greece and the U.K.).

[5] These inconsistent findings between the two studies may be attributed to differences in the specifications used to test the long-run Fisher relationship between stock prices and goods prices. Rather than a bivariate relationship between stock prices and goods prices, Ely and Robinson (1997, p. 148) examine a multivariate relationship that incorporates real output and money as additional variables due to their potential effect on the relationship between stock prices and goods prices.

[6] Luintel and Paudyal (2006) use monthly data for the U.K. retail price index, the Financial Times All-Share Index and seven equally weighted industry portfolios (constructed for consumer goods, financial institutions, general manufacturing, investment trusts, mineral extraction, services and utilities) over the period 1955–2002. In five of the seven industry groups, the price elasticity is greater than 1.0; in one industry (mineral extraction) it is below 1.0; and in one (investment trusts) it is exactly 1.0.

[7] The two-regime TVECM provides a framework for testing the long-run relationship between stock and goods prices, allowing for asymmetric adjustment of stock returns and expected inflation to the long-run equilibrium.

In contrast, cointegration is found in nine countries (Austria, Belgium, France, Germany, Luxembourg, Norway, Spain, Sweden and the U.S.) when Pesaran et al.'s (2001) ARDL bounds test of cointegration is employed. The results show that the long-run Fisher coefficients for all cointegrating vectors are statistically significant and greater than one. The authors conclude that in 9 of 19 countries, investment in stocks can be used as an effective hedge against inflation in the long run. However, the authors could not find any support for the GFH in these 9 countries over both the short and long run when they applied tests based on panel VAR. Overall, these results indicate that the GFH is not a phenomenon that holds globally.

The results are strongly consistent with the GFH reported by Gregoriou and Kontonikas (2010), who employ three panel cointegration techniques developed by Maddala and Wu (1999), Harris and Tzavalis (1999) and Levin et al. (2002). Gregoriou and Kontonikas (2010) test the long-run relationship between stock prices and goods prices for 16 OECD courtiers using monthly data over the period 1970–2006. They conduct tests of the GFH not only for the full sample period but also across three sub-periods (1970–1979, 1980–1989 and 1990–2006) that reflect three distinct inflation regimes: (i) the 1970s period of high inflation, during which the global economy experienced two oil shocks, in 1973 and 1979; (ii) the 1980s period of moderate inflation; and (iii) the later period, during which inflationary pressures were brought under control. Employing the three panel cointegration tests, they find very conclusive evidence that stock and goods prices are cointegrated over all the sample periods. The results based on the Pedroni (2001) fully modified OLS heterogeneous panel cointegration technique show that the panel estimates of the long-run elasticity of stock prices with respect to goods prices are lower than 1.0 in all cases, except for the period during which inflation was brought under control. These elasticity estimates are lower in magnitude than those of other studies, which report coefficients that generally exceed 1.0. The results also show that the null hypothesis of a one-to-one correspondence between stock and goods prices for all periods cannot be rejected, implying that common stocks provide a perfect long-run hedge against inflation. In an attempt to extend the work of Gregoriou and Kontonikas (2010) to pay special attention to the cross-section dependence issue in testing the GFH equation, Omay et al. (2015) employ the more comprehensive panel cointegration methods of Pedroni (2001) (fully modified OLS; FMOLS), Pesaran (2006) and Westerlund (2005) and obtain results lending strong support to the existence of cointegration between stock and goods prices. They also report Fisher coefficient estimates ranging between 0.68 and 1.27, giving support to the GFH.

Bampinas and Panagiotidis (2016) argue that the hedging ability of stocks can be enhanced by constructing investment portfolios of stocks with higher long-run betas as a part of an asset selection and allocation strategy. Using price data for the 25 highest beta stocks that have been continuous constituents of the S&P 500 index over the period 1993–2012, they find that the prices of these stocks tend to move with goods prices in the long run, even though the aggregate index appears to lack long-run inflation hedging ability over the same period. The results indicate that although the aggregate S&P 500 index lacks inflation hedging ability, portfolios constructed from the constituents (stocks) with high long-run beta offer superior hedging ability, regardless of whether the portfolios are constructed on an in-sample or out-of-sample basis.

2.2 Evidence for emerging market countries

A massive amount of empirical work has also been carried out on testing the validity of the GFH for the stock markets of emerging market countries.[8] A number of reasons can be put forward for the increasing attention of researchers to the stock markets of emerging market countries. First, the importance of emerging markets countries has grown dramatically over the last 25 years, since they have produced a superior rate of economic growth compared with major industrialized countries (Hale 2012). Second, the share of the stock markets of emerging market countries in world capitalization increased from 7% in 1990 to 32% in 2009 (European Central Bank 2010). Third, international portfolio diversification strategies are not limited to developed markets but encompass emerging markets as well, and as such emerging markets should have a 20%–40% allocation in a global equity portfolio (Hale 2012). Fourth, a series of crises in the 1990s demonstrated the increasing importance of emerging market countries for the stability of the global financial system (Spyrou 2004).

In fact, historically high average inflation rates, expected returns and volatility of equity stocks have attracted the attention of many researchers to testing whether common stocks in emerging market countries provide investors with a better effective hedge against inflation than those of developed market countries. Barnes et al. (1999) argue that “in high inflation countries – and only in high inflation countries – do nominal return seem to adjust to provide some hedge against inflation”. They examine the performance of the Fisher relation between nominal returns on equity and ‘safe’ assets and inflation rates for 18 developed market and 7 emerging market countries with low-to-moderate and high inflation. Using an unbalanced quarterly data set over the period 1957:02–1996:03,[9] Barnes et al. (1999) find that in only 9 countries – Austria, Chile, Germany, India, Israel, Mexico, New Zealand, Peru, the Philippines and the U.K. – are the coefficients of the contemporaneous inflation rate positive, ranging from 0.13 (the Philippines) to 1.17 (Peru). Nevertheless, the coefficients of the contemporaneous inflation rates are significantly positive only in the 4 countries – Chile, Israel, Mexico and Peru – that experienced the highest rates of inflation over the sample period. Of these countries, Israel has an inflation coefficient that is significantly less than one (0.524), whereas Peru has an inflation coefficient significantly greater than one (1.174). Similar results are obtained by Choudhry (2001), who tests the GFH for four high-inflation countries of Latin America – Argentina, Chile, Mexico and Venezuela – using monthly data over the period 1981:01–1998:12 for the first three countries and 1985:01–1998:06 for Venezuela. Applying regression analysis to the stock return-inflation relation, he shows that common stocks in Argentina and Chile offer a good, if not perfect, hedge against inflation. The coefficients of inflation for Argentina and Chile are positive and significant, with magnitudes close to one, although the restriction $\beta_1 = 1$ is not formally tested. In Mexico and Venezuela, on

[8] The MSCI Emerging Market Index lists 25 countries. Included in these are: Brazil, Chile, China, Colombia, Czech Republic, Egypt, Greece, Hungary, India, Indonesia, Korea, Malaysia, Mexico, Morocco, Qatar, Peru, the Philippines, Poland, Russia, South Africa, South Korea, Taiwan, Thailand, Turkey, and United Arab Emirates.

[9] The sample period varies between 1957:02 and 1996:03 for such countries as Chile (1974:0–1996:03), Germany (1970:01–1996:03), Israel (1980:03–1996:03), Korea (1978:01–1996:03), Mexico (1984:01–1996:03), New Zealand (1961:01–1996:01), Peru (1989:01–1996:03), Portugal (1988:01–1996:03), Sweden (1957:02–1996:04) and Luxemburg (1980:01–1996:03).

the other hand, the relation between nominal return and inflation is non-significant. Spyrou (2004) argues that one possible explanation for why emerging market countries behave differently than industrialized countries with respect to the inflation-stock return relation is that money rather than real activity caused inflation in these countries in the 1990s. Examining the inflation-stock return relation for 10 emerging market countries – Argentina, Brazil, Chile, Hong Kong, Korea, Malaysia, Mexico, the Philippines, Thailand and Turkey – using monthly data over the period 1989:01–2000:08, he finds results showing that the relation between stock returns and inflation is significantly positive for Argentina, Malaysia and the Philippines, while it is insignificantly positive for Brazil, Korea, and Mexico. In contrast, the underlying relationship turns out to be insignificantly negative for all the remaining countries, except Thailand.

Alagidede (2009) tests the GFH for six African countries – Egypt, Kenya, Morocco, Nigeria, South Africa and Tunisia – by regressing monthly stock returns on inflation rates over a longer investment horizon of 60 months and finds a positive relation between stock returns and inflation rates for Kenya and Nigeria. He also applies an instrumental variable method and obtains consistent results, thus confirming the validity of the GFH for three markets: Kenya and Nigeria at a 12-month horizon and Tunisia at a 60-month horizon. These results suggest that investors expect stocks to be a good hedge against inflation over long horizons. Bekaert and Wang (2010) also obtain results that lend strong support to the proposition that common stocks offer better inflation hedging in emerging markets than in developed markets. They examine the inflation hedging ability of stocks (and other assets) and provide estimates of 'inflation betas' for over 48 developed and emerging market countries. Using monthly data on stock returns and inflation rates over the period January 1970–January 2010,[10] they obtain results from 'pooled regressions' indicating that OLS estimates of the inflation betas for the equity stocks of all developed markets are insignificantly negative, whereas those for all emerging markets are significantly positive and close to unity. The results also show that for stock returns, the developed market betas increase with horizon but remain significantly below 1, even at the 5-year horizon, whereas those of emerging markets show little horizon dependence and remain at approximately 1.

Al-Khazali and Pyun (2004) and Alagidede and Panagiotidis (2010) apply cointegration techniques to test the GFH on the basis of the model specification using level data on stock and goods prices to examine whether investments in common stocks of emerging market countries provide a better hedge against inflation in the long run. Al-Khazali and Pyun (2004) test the validity of the GFH in nine Pacific-Basin equity markets – Australia, Hong Kong, Indonesia, Japan, South Korea, Malaysia, the Philippines, Singapore, and Thailand – that have grown increasingly important to the safety and vitality of global financial markets. Since the Asian financial crisis of 1997–1998, which had negative ripple effects on the U.S. and European markets, these Pacific-Basin countries have become important for investors and regulators, who want to know whether common stocks in these emerging markets provide better hedges against inflation. Regression analysis shows that nominal stock returns and inflation rates are negatively related in these nine markets in the short run, whereas cointegration analysis based on Johansen's (1988) method confirms a positive relationship between the same variable

[10] While the sample period for most of the countries starts with January 2010, the starting point varies from country to country, between January 1970 and January 2005.

over the long run. Using monthly data over the period 1980:01–2001:12, it is shown that the long-run relation between stock and goods prices, as predicted by the GFH, has a stock price elasticity that ranges from 1.02 (the Philippines) to 1.67 (Hong Kong). Overall, the Fisher effect coefficients are significantly greater than unity in all cases, except Indonesia, the Philippines and Thailand. Based on the impulse response function, it is also shown that the time path of the response of stock prices to a shock in goods prices exhibits an initial negative response that turns positive over the long run. The evidence from the error-correction model shows that the estimated coefficient of speed of adjustment is between 0.01 and 0.03, which means that stock prices take a long time to return to their long-run relation following unexpected movements in goods prices. These results are interpreted as reconciling the conflicting evidence on the inverted short-run and the positive long-run Fisher effects reported in the literature. Similar results are reported by Alagidede and Panagiotidis (2010), who test the long-run relationship between stock and goods prices for six emerging market countries in Africa – Egypt, Kenya, Morocco, Nigeria, South Africa and Tunisia – using an unbalanced dataset over the period 1990–2007. These researchers explore whether equity markets protect investors against inflation in African countries that have adopted inflation targeting (South Africa) and those with high inflation rates. Employing parametric and non-parametric cointegration techniques and the impulse response function, they find support for a positive long-run relationship between stock and goods prices, with a point estimate of the elasticity of stock prices with respect to goods prices ranging from 0.015 (Tunisia) to 2.264 (South Africa). The authors confirm that the estimated speed of adjustment coefficient ranges from 0.0013 (South Africa) to 0.155 (Tunisia); hence, it takes a long time for stock prices to return to their equilibrium following a short-run deviation. They also find that the time path of the response of stock prices to a shock in goods prices exhibits a transitory negative response in Egypt and South Africa that becomes positive over longer horizons. This indicates that common stocks tend to provide a hedge against rising goods prices in African markets.

3 Data, model and method

3.1 Data

The dataset examined in this paper comprises monthly observations of stock market and consumer price[11] indexes for 50 emerging markets covering four continents and several regions, namely, South America (Brazil, Chile, Colombia, Mexico, Peru and Venezuela), Africa (Ivory Coast, Morocco, Nigeria, South Africa and Tunisia), East Asia (Hong Kong, Korea, Taiwan and China), South Asia (Bangladesh, India, Pakistan and Sri Lanka), Southeast Asia (Indonesia, Malaysia, the Philippines, Singapore, Thailand and Vietnam), the Middle East (Bahrain, Egypt, Israel, Jordan, Kuwait, Oman, Qatar, Saudi Arabia, United Arab Emirates and Turkey), Central Asia (Kazakhstan)[12] and Eastern and Central Europe (Bulgaria, Croatia, Czech Republic, Estonia, Greece, Hungary, Latvia, Lithuania,

[11] One reason for using monthly data is that the majority of studies utilized the monthly frequency of data on stock and goods prices. Another reason is to increase the number of sample observations for employing cointegration analysis and to make our results comparable with those of other studies in the underlying area.

[12] Due to the unavailability of data on their stock markets, other countries from Central Asia (e.g., Kyrgyzstan, Tajikistan and Uzbekistan) are not included in the dataset.

Poland, Romania, Slovak Republic, Slovenia, Ukraine, and Russia). The sample is not balanced because of either insufficient or unreliable CPI and stock market data. To ensure sufficient statistical power of the tests employed under the cointegration framework, only those markets for which at least 5 years of uninterrupted monthly observations on CPI and stock market prices are available are included in this study. The details are shown in Table 1, in which the sample for each country begins with the month of a year for which CPI and stock market data are available and ends in 2016 M01. The data were obtained from Datastream.

3.2 Model and method

The proposition that common stocks should provide a perfect hedge against inflation is embedded in the GFH.[13] Associated with perfectly competitive and informationally efficient capital markets in which investors are rational, this hypothesis postulates that stock prices should move one-for-one with goods prices to compensate investors for rising prices, as common stocks are claims against real assets (see, for example, Bodie 1976) and, consequently, a one-period expected nominal stock return should adjust fully to the expected inflation over the holding period such that the expected real stock return remains not only constant but also independent of expected inflation over time.[14] This implies that two alternative specifications can be used to test the validity of the GFH: one based on level data on stock and goods prices and another based on first-difference data on nominal stock returns and inflation. The model specification using level data on stock and goods prices can be represented in a general stochastic regression form as follows:

$$s_t = \gamma + \theta p_t + u_t \tag{1}$$

where $s_t(p_t)$ is the logarithm of the stock price (consumer price) index representing a weighted sum of the prices of common stocks (commodities) included in the market portfolio (consumer basket), γ and θ are the coefficients of the Fisher relation, and u_t measures the deviations of s_t corresponding to each level of p_t. Thus, if the GFH holds precisely and common stocks provide a perfect hedge against inflation over time, then the restriction that $\theta = 1$ should not be rejected. If $\theta < 1$, then common stocks should offer a partial hedge against inflation, whereas if $\theta > 1$ then common stocks should rather offer a superior hedge.

[13] Arguably, Fisher (1930) was the first economist to develop the standard hypothesis relating the nominal interest rate to expected inflation and real interest rates. However, Humphrey (1983, p. 2-6) argues that the proposition that the nominal interest rate equals the real interest rate plus the expected inflation rate has a longer history that dates back more than 240 years to the writings of William Douglass, Henry Thornton, John Stuart Mill, Jacob de Hass, Alfred Marshall and J.B. Clark. Moreover, it has been argued that the claim is disproved by Fisher (1896) himself when he noted that he was by no means the first to present that analysis.

[14] When applied to Treasury bills and deposit markets (see, inter alia, Fama and Schwert 1977; Rose 1988; MacDonald and Murphy 1989; Mishkin 1992; Balparda et al. 2016) the FH implies that the one-period nominal interest rate on a Treasury bill (a bank deposit) should fully reflect the inflation rate anticipated by market agents such that the *ex ante* real rate of return remains constant and independent of anticipated inflation over time. This may be defined as follows:

$$i_t = r^e_{t+1} + \Delta p^e_{t+1}$$

Table 1 List of countries and sample period

Country	Sample period	
Africa		
Ivory Coast	1998 M09	2016 M01
Morocco	2002 M01	2016 M01
Nigeria	2000 M01	2016 M01
South Africa	1995 M06	2016 M01
Tunisia	1997 M12	2016 M01
East Asia		
Hong Kong	1980 M10	2016 M01
Korea	1983 M05	2016 M01
Taiwan	1980 M01	2016 M01
China (mainland)	1990 M12	2016 M01
South Asia		
Bangladesh	2004 M01	2016 M01
India	1980 M01	2016 M01
Pakistan	1994 M05	2016 M01
Sri Lanka	1993 M06	2016 M01
South East Asia		
Indonesia	1990 M04	2016 M01
Malaysia	1982 M01	2016 M01
Philippines	1987 M01	2016 M01
Singapore	1999 M08	2016 M01
Thailand	1982 M01	2016 M01
Vietnam	2000 M07	2016 M01
Middle East		
Bahrain	2007 M08	2016 M01
Egypt	1998 M01	2016 M01
Israel	1992 M10	2016 M01
Jordan	1996 M01	2016 M01
Kuwait	1997 M03	2016 M01
Oman	2001 M01	2016 M01
Qatar	2009 M01	2016 M01
Saudi Arabia	1998 M10	2016 M01
United Arab Emirates	2007 M01	2016 M01
Turkey	1988 M01	2016 M01
East Europe		
Baltic states		
Estonia	1996 M06	2016 M01
Latvia	2000 M01	2016 M01
Lithuania	2000 M01	2016 M01
Former Soviet states		
Kazakhstan	2000 M07	2016 M01
Russian Federation	1995 M09	2016 M01

Table 1 (continued)

Country	Sample period	
Ukraine	1997 M10	2016 M01
Central Europe		
Croatia	1998 M01	2016 M01
Czech Republic	1993 M09	2016 M01
Hungary	1991 M01	2016 M01
Poland	1991 M04	2016 M01
Slovak Republic	1993 M09	2016 M01
Slovenia	2006 M04	2016 M01
Southeastern Europe		
Bulgaria	2000 M10	2016 M01
Greece	1991 M01	2016 M01
Romania	1997 M09	2016 M01
South America		
Brazil	1982 M01	2016 M01
Chile	1993 M09	2016 M01
Colombia	2001 M07	2016 M01
Mexico	1987 M01	2016 M01
Peru	1992 M01	2016 M01
Venezuela	2007 M12	2016 M01

In an alternative specification using first-difference data on stock and goods prices (nominal stock return and inflation), the GFH suggests that the one-period expected nominal stock return should be equal to the sum of the expected real stock return plus the expected inflation, which can be approximated by:

$$\Delta s^e_{t+1} = r^e_{t+1} + \Delta p^e_{t+1} \tag{2}$$

where $\Delta s^e_{t+1}\left(\Delta p^e_{t+1}\right)$ is the logarithm of the expected rate of change in stock (goods) prices, that is the expected nominal return on the stock market portfolio (the basket of consumer goods) at time t + 1, and r^e_{t+1} is the expected real stock return. To make eq. (2) empirically testable, let us assume that the expected real return, r^e_{t+1}, is equal to a constant value, α, over time plus a random term, v_t. Thus, the behavior of the expected real return on common stocks can formally be modeled as follows:

$$r^e_{t+1} = a + v_t \tag{3}$$

In addition, let us assume that if economic agents are rational and if the available information set, Ω_t, is processed efficiently to predict future changes in stock and goods prices, then *ex post* changes in stock and goods prices realized from time *t* to *t + 1* will differ from *ex ante* changes by mean zero serially uncorrelated random terms, that is $E(u_{1t+1}|\Omega_t) = 0$; $E(u_{2t+1}|\Omega_t) = 0$; $E(u_{1t+1}u_{1t+1-i}) = 0\ \forall_i \neq 0$; $E(u_{2t+1}u_{2t+1-i}) = 0\ \forall_i \neq 0$. Formally,

$$\Delta s_{t+1} = \Delta s^e_{t+1} + u_{1t+1} \tag{4}$$

$$\Delta p_{t+1} = \Delta p^e_{t+1} + u_{2t+1} \tag{5}$$

By substituting eq. (2) into eqs. (3) to (5) and rewriting the resulting expression in a stochastic regression form, we obtain:

$$\Delta s_{t+1} = \alpha + \beta \Delta p_{t+1} + \varepsilon_{t+1} \tag{6}$$

where $\alpha = a$ and $\varepsilon_{t+1} = v_t + u_{1t+1} - u_{2t+1}$ is the composite error term, representing the effect on stock market returns that is unexplained by expected inflation. In a contemporaneous stochastic regression form, eq. (6) can be rewritten as follows:

$$\Delta s_t = \alpha + \beta \Delta p_t + \varepsilon_t \tag{7}$$

For the GFH to hold precisely and for common stocks to provide a perfect hedge against expected and contemporaneous inflation, the restriction $(\alpha, \beta) = (0, 1)$ should not be rejected, and $\varepsilon_{t+1}(\varepsilon_t)$ should be white noise.[15]

It is argued that the GFH based on the specification using first-difference data on stock and goods prices essentially represents a short-run rather than a long-run relation.[16] Unlike the specification based on first-difference data, the specification based on level data is more appealing because it is conducive to employing cointegration analysis. The advantage of this procedure is that it has been demonstrated by Stock (1987) that if the underlying series are cointegrated, then OLS will produce consistent (or super-consistent) estimates of the regression parameters, despite such problems as simultaneity, autocorrelation and heteroskedasticity. Second, as Hendry (1986, p. 54) demonstrates, when a time series is first differenced, the long-run information contained in the level data is lost. It follows that the long-run information can be utilized fully by testing the GFH using the specification based on level data. Third, as demonstrated by Granger (1986) and Engle and Granger (1987), first-differenced variables that are I(0) have only limited memories of past behavior, whereas variables I(1) have infinitely long memories (i.e., an innovation will permanently affect the process). Fourth, a first-differenced data model is unlikely to perform well, as has been demonstrated by the majority of empirical studies. Finally, as demonstrated by Stock (1987) and Park and Phillips (1988), a level data model produces estimates converging on the true parameters at the rate N rather than $\sqrt{N}$ (where N is the sample size).

For a pair of variables, s_t and p_t, underlying eq. (1) to form a cointegrating (long-run) relation, a necessary condition is that both the variables are integrated at the same order

[15] The OLS estimate of the slope coefficient of the Fisher relation (β) may be greater than 1.0, as demonstrated by Darby (1975), or less than 1.0, as demonstrated by Mundell (1963) and Tobin (1965). Consequently, if $\beta \geq 1.0$ ($\beta \leq 1.0$), then common stocks may provide a superior (partial) hedge against inflation. For the GFH to hold precisely, the expected (current) real stock returns should also be independent of the expected (current) inflation rates. The relationship between real stock returns (rsr_t) and inflation rates (Δp_t) can be obtained by substituting the Fisher equation ($\Delta s_t = rsr_t + \Delta p_t$) in equation (7) and rearranging the resulting expression as follows: $rsr_t = \alpha + b\Delta p_t + \varepsilon_t$, where $b = (1 - \beta)$ should be equal to zero if GFH holds (see, for example, Rushdi et al. 2012).

[16] See, for example, Anari and Kolari (2001), Al-Khazali and Pyun (2004) and Alagidede and Panagiotidis (2010).

of unity (i.e., s_t~$I(1)$ and p_t~$I(1)$), and a sufficient condition requires the linear combination thereof to be integrated at order zero (i.e., u_t~$I(0)$). It must, however, be noted that cointegration is a necessary but not sufficient condition for the long-run inflation-hedging ability of common stocks to exist. The sufficient condition requires the restriction $\theta = 1$ to be satisfied. In addition, if s_t and p_t are cointegrated in the long run, then a valid error-correction representation must exist between them that can be normalized on both s_t and p_t as follows:

$$\Delta s_t = \varphi_1 + \sum_{i=1}^{k} \delta_{1i} \Delta p_{t-i} + \sum_{i=1}^{k} \phi_{1i} \Delta s_{t-i} + \eta_1 (s_{t-1} - \gamma - \theta p_{t-1}) + e_{1t} \tag{8a}$$

$$\Delta p_t = \varphi_2 + \sum_{i=1}^{k} \delta_{2i} \Delta p_{t-i} + \sum_{i=1}^{k} \phi_{2i} \Delta s_{t-i} + \eta_2 (s_{t-1} - \gamma - \theta p_{t-1}) + e_{2t} \tag{8b}$$

where $\eta_1(\eta_2)$ is the coefficient of the error-correction term normalized on stock prices (goods prices) that measures the speed of adjustment towards the long-run equilibrium relation.

Following Anari and Kolari (2001), Al-Khazali and Pyun (2004) and Alagidede and Panagiotidis (2010), we test the long-run relation between stock and goods prices by employing Johansen's (1988) maximum likelihood cointegration test, which is considered more powerful than residual-based cointegration tests on the following grounds. First, its results are invariant with respect to the direction of normalization. Second, it provides estimates of all cointegrating vectors existing within a system of variables. Third, it allows us to test a priori restrictions on the coefficients of the cointegrating vectors imposed by economic theory. Finally, it fully captures the underlying time series properties of data. The Johansen test is based on a multivariate VAR of n variables given by

$$Z_t = \mu + \sum_{r=1}^{k} \Pi_r Z_{t-r} + \upsilon_t, \tag{9}$$

where Z_t is an $n \times 1$ vector of I (1) variables, $\Pi_1, \Pi_2, \ldots, \Pi_k$ are matrices of unknown parameters, μ is a vector of constants, υ_t is a vector of Gaussian error terms and k is the maximum lag or the order of the VAR. this model can be reparametrized as follows:

$$\Delta Z_t = \mu + \sum_{r=1}^{k-1} \Gamma_r \Delta Z_{t-r} + \Gamma_k Z_{t-k} + \upsilon_t \tag{10}$$

where $\Gamma = -I + \Pi_1 + \Pi_2 + \ldots + \Pi_k$ and I is the identity matrix. Π is the cointegrating matrix such that $\Pi Z_t = 0$ represents the long-run equilibrium. An important question that needs to be explained here concerns the rank, r, of the matrix Π, and there are three possibilities. If Π has a full rank matrix (i. e. , $r = n$), any linear combination of Z_t variables will be stationary. In the two-variable model represented by equation (1), this can only occur if s_t and p_t are stationary, which means that the correct model will be in level rather than first difference. If, on the other hand, Π has a zero rank matrix (i. e. , $r = 0$), then any linear combination of Z_t variables will be nonstationary, which means that the variables are not cointegrated and the proper model will be in first difference. Finally, If Π has a less than full rank matrix ($0 < r < n$) then we can write

$$\Gamma = \alpha\beta' \tag{11}$$

where the columns of the $n \times r$ matrix β are the cointegrating vectors and the columns of the $n \times r$ matrix α are error-correction coefficients measuring the speed of convergence of the dependent variable towards the long-run equilibrium state. The following are important steps involved in the application of the Johansen technique of cointegration.

(1) Regress ΔZ_t and Z_{t-k} on the lagged differences of ΔZ_t and a constant as follows:

$$\Delta Z_t = \mu + \sum_{r=1}^{k-1} \Gamma_{0r} \ \Delta Z_{t-r} + V_{0t} \tag{12}$$

$$Z_{t-k} = \mu + \sum_{r=1}^{k-1} \Gamma_{1r} \ \Delta Z_{t-r} + V_{1t} \tag{13}$$

(2) Extract the residuals V_{0t} from equation (12) and V_{1t} from equation (13).
(3) Use the residuals V_{0t} and V_{1t} to calculate n-squared canonical correlations (or the eigenvalues) with order $\lambda_1 > \lambda_2 > \lambda_3 > \ldots > \lambda_n$.
(4) Use the eigenvalues to construct two test statistics for testing the existence of the number of unique cointegrating vectors between Z_t variables. The first statistic, known as the maximum eigenvalue (*Max*) test, evaluates the null hypothesis that there are exactly r cointegrating vectors and is given by

$$Max = -T \ ln(1-\lambda_{r+1}) \tag{14}$$

where T is the sample size. The second statistic, known as the *Trace* test, evaluates the null hypothesis that there are at most r cointegrating vectors and is given by

$$Trace = -T \ \sum_{t=r+1}^{n} \ln(1-\lambda_t) \tag{15}$$

where $\lambda_{r+1}, \lambda_{r+2}, \ldots, \lambda_n$ is the $n-r$ smallest squared canonical correlations of V_{0t} with respect to V_{1t}. Johansen and Juselius (1990) note that the power of the *Trace* test is lower than that of the *Max* test. In both the cases the null hypothesis is rejected if the calculated value of the test statistic is greater than the critical value as tabulated in Johansen and Juselius (1990).

4 Empirical results

4.1 Unit root results

Before testing for a long-run relationship between stock and goods prices, unit root tests are conducted first to determine the order of integration of the underlying series. To this end, two unit root tests are employed: the Dickey and Fuller (1979) and Phillips and Perron (1988) tests. In the presence of any conflict in the unit root results, the judgement of whether a particular series is stationary or non-stationary will be based on the latter test, since it is considered to be more robust with respect to a wide variety of serial correlation and time-dependent heteroscedasticity.[17] While three different forms of a regression model can be used depending on whether or not the model from which the underlying test statistics are calculated includes a constant term and a time trend, we use the model regressing the first difference (level) of a series on its lagged level plus a constant term when applying the Dickey-Fuller (Phillips-Perron) test to determine if the underlying series is difference stationary rather than trend stationary.[18] The results of the Dickey-Fuller (ADF)[19] and Phillips-Perron (PP) tests, as reported in Table 2,[20] are consistent in showing that only in 26 of the 50 countries are the stock and goods prices I(1) in level and I(0) in first difference. In the two cases of Saudi Arabia and Latvia, stock prices are I(1) when both the ADF and PP tests are used, whereas goods prices are I(1) when only the latter test is used. This incompatibility in the unit root results of the ADF and PP tests can be dealt with by relying on the outcome of the latter test, which suggests that both goods and stock prices are I(1).

That notwithstanding, in the remaining 22 countries, the order of integration of the underlying series of stock and goods prices is predominantly incompatible when both the ADF and PP tests are used with only one exception, i.e., Mexico, where both tests consistently indicate that stock and goods prices are I(0)). The results show that stock prices turn out to be I(1) in level for 16 countries (Czech Republic, Estonia, Greece, Hong Kong, Hungry, Israel, Korea, the Philippines, Poland, Romania, Russia, Slovak Republic, Slovenia, Taiwan, Turkey and Venezuela) and I(0) for 5 countries (Brazil, Chile, China, Columbia and Mexico) when both the ADF and PP tests are used. In the case of Peru, stock prices are I(1) when ADF is used. However, relying on the PP test, stock prices turn out to be I(0).

Only in 10 countries (Estonia, Hungry, Israel, Korea, Mexico, Peru, the Philippines, Poland, Romania and Taiwan) are goods prices I(0) when the ADF and PP tests are used. In contrast, there are 7 countries (Brazil, Czech Republic, Hong Kong, Greece, Slovak Republic, Slovenia and Turkey) in which case goods prices turn out to be I(2) in level when adjudged on the basis of the ADF test but I(0) when tested on the basis of the PP test. In the case of Russia, the ADF test supports an I(1) process, whereas the PP results indicate stationarity. For Venezuela, the results show that goods prices are consistently I(2) when both the ADF and the PP are used.

Thus, there are 18 countries (Chile, China, Columbia, Czech Republic, Estonia, Greece, Hong Kong, Hungry, Israel, Korea, the Philippines, Poland, Romania, Russia,

[17] Moosa and Bhatti (1997, p. 149) argue, "If normality, serial correlation or heteroscedasticity statistics are significant, the Phillips-Perron procedure should be adopted."

[18] The time trend can be included in the regression model to allow the alternative of trend stationarity. In all cases, the null of hypothesis of non-stationarity (unit root) is rejected when the test statistic is significantly negative and greater in absolute terms than the critical value at the appropriate significance level. Because the PP test statistics are asymptotically equivalent to the corresponding Dickey-Fuller statistics, the critical values to be used in conjunction with the Phillips-Perron test are identical to those tabulated in Fuller (1976) and MacKinnon (1996). Both the test statistics calculated from the regression model with a time trend is valid (in the sense that it accurately confirms or refutes stationarity) only if the coefficient of the time trend is zero, but if there is a significant time trend, then the underlying series will not be stationary even if the coefficient of the lagged value of the series is less than zero. See, for example, Moosa and Bhatti (1997, p. 144-151) for a detailed discussion of these unit root tests.

[19] The optimal number of lags in the ADF test is determined based on the Schwartz Information Criterion (SIC).

[20] Tests will be not applied to the first difference of a series if it turns out to be I(0) in level. This is why the outcomes for level and first difference are reported for those series which are I(1) in level.

Slovak Republic, Slovenia, Taiwan and Turkey) for which stock and goods prices are either I(1) or I(0) when the judgement is based on the PP test. Moreover, there are three countries (Brazil, Mexico and Peru) in which both stock and goods prices turn out to be I(0) when the judgement is based on the PP test; in addition, there is the unique case of Venezuela, where the ADF and PP tests are consistent in showing that stock and goods prices are I(1) and I(2), respectively.

That stock prices are I(0) for Brazil, Chile, China, Columbia and Mexico may be a quite surprising outcome, since the majority of the literature looking at stock market performance has reported that stock prices are traditionally an I(1) process. However, this inconsistent evidence on stock prices is not a new phenomenon. There is some evidence in the literature to suggest that stock prices are I(0). For example, Ely and Robinson (1997, p. 146) and Hassan et al. (2015, p. 140) find that when examined on the basis of the ADF test, the stock-price index levels for Germany and Canada are I(0), respectively. In contrast, there is some evidence in the literature to suggest that consumer prices may be I(2) and I(0). For example, Adrangi et al. (1999, p. 67) and Hassan et al. (2015, p. 140) find on the basis of the ADF test that consumer prices are I(0) for Peru, Finland and Ireland, whereas Phylaktis and Blake (1993), among others, find consumer prices to be I(2) for three high-inflation countries (Argentina, Brazil and Mexico) and Mahdavi and Zhou (1994) find the ratio of the domestic to the foreign price index to be I(2)for several countries, including Brazil, Israel, Mexico, Peru and Turkey.

4.2 Cointegration results

Prior to testing the GFH in the long run using the model specification based on level data on stock and goods prices, we use the model specification based on first-difference data on the underlying variables, as represented by equation (7), to see whether there is a significantly positive relationship between nominal stock return and inflation and whether the GFH fares well in the short run. The results (available upon request) show that the OLS estimates of the slope coefficients of the short-run Fisher relation are predominantly positive (in 29 of the 50 cases). However, only in a few cases are the coefficients significantly positive. The Fisher coefficients are significant at the 5% level for Brazil and Jordan, whereas for Peru, Turkey and Kazakhstan, the coefficients are marginally significant at the 10% level. In contrast, for Saudi Arabia, Russia and Latvia, the coefficients are significantly negative at the 5% level. Furthermore, the Wald test of the restrictions that $[\beta_1 = 1]$ and $[\beta_0 = 0, \beta_1 = 1]$ indicates that the GFH holds precisely only in 4 countries (Jordan, Turkey, Kazakhstan and Peru), implying that stocks are a perfect hedge against inflation.

The empirical testing of the GFH based on the model specification using the level data on stock and goods prices is conducted based on Johansen's (1988) and Johansen and Juselius's (1990) maximum likelihood method of cointegration for only 28 countries for which the order of integration of stock and goods prices is the same based on either the ADF and PP tests or the PP test alone (for Saudi Arabia and Latvia).[21] In

[21] An important step in the application of this technique is fixing the optimal lag level. The lag level was set to 4, which in almost all cases removes serial correlation in the residuals of the VAR model. Moreover, an unrestricted constant specification is used to account for the trending behavior of the underlying series (for more details on the specification of deterministic terms, see Johansen and Juselius (1990) and Johansen (1995)).

Table 2 Unit root tests

Country	Intercept			
	ADF		PP	
	Levels	Diff	Levels	Diff
Africa				
Ivory Coast				
S_t	0.22	−12.80*	−0.01	−12.84*
P_t	−1.19	−12.43*	−1.16	−12.40*
Morocco				
S_t	−1.69	−11.77*	−1.63	−12.04*
P_t	−0.51	−11.55*	−1.13	−12.06*
Nigeria				
S_t	−2.72	−11.51*	−2.59	−11.56*
P_t	−2.04	−11.97*	−2.45	−11.95*
South Africa				
S_t	−0.66	−16.21*	−0.66	−16.21*
P_t	−0.88	−11.30*	−0.89	−11.51*
Tunisia				
S_t	−0.43	−13.49*	−0.47	−13.52*
P_t	2.65	−10.96*	3.07	−10.93*
East Asia				
Hong Kong				
S_t	−1.27	−19.38*	−1.26	−19.39*
P_t	−1.19	−2.58	−4.44*	
Korea				
S_t	−2.16	−18.05*	−2.15	−18.05*
P_t	−3.04*		−3.01*	
Taiwan				
S_t	−1.99	−18.96*	−2.00	−18.99*
P_t	−3.64*		−3.81*	
China (Mainland)				
S_t	−3.30*		−3.30*	
P_t	−1.71	−5.75*	−1.88	−14.56*
South Asia				
Bangladesh				
S_t	−2.64	−11.70*	−2.62	−11.73*
P_t	−0.93	−3.39*	−0.15	−7.54*
India				
S_t	−1.49	−19.08*	−1.47	−19.08*
P_t	−0.97	−4.10*	−1.60	−13.15*
Pakistan				
S_t	0.18	−15.80*	0.17	−15.80*

Table 2 (continued)

Country	Intercept			
	ADF		PP	
	Levels	Diff	Levels	Diff
P_t	−0.01	−6.01*	−0.06	−13.13*
Sri Lanka				
S_t	−0.25	−14.32*	−0.45	−14.42*
P_t	−1.27	−6.78*	−1.32	−11.17*
Southeast Asia				
Indonesia				
S_t	−0.26	−14.30*	−0.15	−14.30*
P_t	−1.46	−8.67*	−1.43	−9.13*
Malaysia				
S_t	−1.70	−11.36*	−1.33	−17.80*
P_t	−0.59	−15.90*	−0.67	−15.54*
Philippines				
S_t	−1.54	−16.39*	−1.60	−16.28*
P_t	−5.79*		−6.68*	
Singapore				
S_t	−1.59	−12.21*	−1.67	−12.37*
P_t	0.00	−5.89*	0.30	−16.21*
Thailand				
S_t	−1.89	−18.16*	−1.93	−18.15*
P_t	−1.63	−14.79*	−1.72	−14.78*
Vietnam				
S_t	−2.67	−9.13*	−2.65	−8.93*
P_t	−0.16	−7.16*	−0.04	−7.38*
Middle East				
Bahrain				
S_t	−1.90	−5.76*	−1.78	−5.90*
P_t	−0.51	−11.40*	−0.19	−12.21*
Egypt				
S_t	−0.95	−12.33*	−1.04	−12.74*
P_t	2.00	−9.98*	2.69	−10.02*
Israel				
S_t	−1.08	−15.03*	−1.10	−15.06*
P_t	−5.51*		−7.13*	
Jordon				
S_t	−1.44	−8.14*	−1.32	−12.73*
P_t	0.05	−12.64*	0.01	−12.55*
Kuwait				
S_t	−1.14	−9.26*	−1.20	−9.39*
P_t	1.81	−17.64*	1.68	−17.45*

Table 2 (continued)

Country	Intercept			
	ADF		PP	
	Levels	Diff	Levels	Diff
Oman				
S_t	−2.24	−5.59*	−1.84	−10.31*
P_t	−0.52	−4.28*	−0.20	−10.96*
Qatar				
S_t	−2.25	−10.11*	−2.27	−10.05*
P_t	1.88	−6.41*	1.12	−6.36*
Saudi Arabia				
S_t	−2.12	−11.60*	−2.03	−11.83*
P_t	0.92	−2.58	2.09	−11.23*
United Arab Emirates				
S_t	−1.88	−7.99*	−1.58	−8.06*
P_t	−2.17	−4.62*	−1.80	−9.26*
Turkey				
S_t	−2.17	−17.32*	−2.12	−17.35*
P_t	−2.40	−1.12	−6.60*	
Eastern Europe				
Baltic states				
Estonia				
S_t	−1.72	−11.90*	−1.70	−11.88*
P_t	−3.03*		−2.89*	
Latvia				
S_t	−2.38	−11.73*	−2.22	−11.75*
P_t	−1.80	−1.83	−1.29	−9.75*
Lithuania				
S_t	−1.40	−9.82*	−1.39	−9.91*
P_t	−0.38	−9.81*	−0.35	−9.83*
Former Soviet states				
Kazakhstan				
S_t	−1.61	−8.83*	−1.60	−8.87*
P_t	0.18	−7.36*	0.22	−5.79*
Russian Federation				
S_t	−2.16	−12.60*	−1.98	−12.69*
P_t	−2.35	−5.68*	−3.24*	
Ukraine				
S_t	−1.20	−10.74*	−1.22	−11.01*
P_t	−0.28	−6.90*	−0.13	−6.66*
Central Europe				
Croatia				
S_t	−1.45	−13.26*	−1.59	−13.35*

Table 2 (continued)

Country	Intercept			
	ADF		PP	
	Levels	Diff	Levels	Diff
P_t	−2.42	−10.30*	−2.61	−10.29*
Czech Republic				
S_t	−1.97	−12.85*	−2.23	−12.69*
P_t	−2.77	−2.33	−6.21*	
Hungary				
S_t	−1.49	−15.39*	−1.48	−15.41*
P_t	−2.93*		−10.64*	
Poland				
S_t	−2.85	−15.63*	−2.64	−16.34*
P_t	−3.17*		−19.65*	
Slovak Republic				
S_t	−1.55	−11.87*	−1.73	−11.49*
P_t	−2.23	−2.26	−5.26*	
Slovenia				
S_t	−1.28	−7.40*	−1.16	−7.79*
P_t	−2.68	−1.26	−3.46*	
Southeastern Europe				
Bulgaria				
S_t	−1.96	−9.89*	−1.95	−10.17*
P_t	−1.85	−9.64*	−2.05	−9.72*
Greece				
S_t	−1.03	−15.93*	−1.22	−15.99*
P_t	−2.27	−2.79	−8.21*	
Romania				
S_t	−0.88	−12.06*	−1.01	−12.10*
P_t	−8.43*		−15.84*	
South America				
Brazil				
S_t	−4.18*		−3.69*	
P_t	−2.71	−2.06	−3.44*	
Chile				
S_t	−3.12*		−3.12*	
P_t	2.35	−8.55*	2.45	−8.64*
Colombia				
S_t	−3.19*		−3.00*	
P_t	−1.47	−6.35*	−1.91	−5.76*
Mexico				
S_t	−3.46*		−3.45*	
P_t	−4.15*		−6.58*	

Table 2 (continued)

Country	Intercept			
	ADF		PP	
	Levels	Diff	Levels	Diff
Peru				
S_t	−2.36	−9.17*	−2.91*	
P_t	−3.40*		−9.24*	
Venezuela				
S_t	1.92	−7.53*	1.55	−7.53*
P_t	3.04	−1.86	6.41	−1.73

ADF = Augmented Dickey-Fuller; PP=Philip-Perron; S_t is the stock market index; P_t is the consumer price index. *p*-values (available upon request) for ADF and PP tests are from MacKinnon (1996); * denotes significance at the 5% level. Lag lengths are based on the Schwartz information criterion (SIC)

contrast, for 18 countries for which the order of integration is not the same (one series is I(0) and another series is I(1)), we employ Pesaran et al.'s (2001) ARDL bounds test of cointegration, which does not require the same order of integration of the underlying series. There are 3 cases (Brazil, Mexico and Peru) for which regression analysis can be employed, since both the underlying series are I(0) when the PP test is used. There is only one case, Venezuela, for which neither OLS nor cointegration is valid.

The results of the Johansen test, as reported in Table 3, show that the null hypothesis of no cointegration between stock and goods prices is rejected on the basis of the *Max* and *Trace* tests in only 10 cases. Moreover, in 6 of these 10 cases (South Africa, Singapore, Egypt, Oman, Latvia and Lithuania), the null hypothesis of no cointegration can be rejected at the 5% significance level, whereas in 4 cases (Malaysia, Thailand, Vietnam and Kuwait), it can be rejected at the 10% significance level. The results suggest the presence of two cointegrating vectors in the case of three countries – Thailand, Oman and Latvia – although the null hypothesis that there is at least one cointegrating vector was rejected at the 10% level, except for Latvia.

The estimated values of the coefficient of the GFH, as reported in Table 4, are correctly signed in all the cases and range from 0.51 (Egypt) to 3.29 (Lithuania). The results (as shown in Column 6 of Table 4) show that the restriction on the coefficient of the GFH that $\theta = 1$ holds in two cases (Oman and Vietnam), whereas it does not in six cases (Latvia, Lithuania, Malaysia Singapore, South Africa and Thailand) for which the estimated value of the underlying Fisher coefficient is significantly greater than one (i.e., $\theta > 0$). In contrast, the estimated values of the Fisher coefficient are less than one in two cases (Egypt and Kuwait), albeit they are insignificant. Thus, common stocks clearly provide a superior hedge of inflation in six cases (South Africa, Singapore, Malaysia, Lithuania, Latvia and Thailand) in which the estimated values of the Fisher coefficient θ are greater than one, whereas they provide a perfect hedge of inflation in two cases (Oman and Vietnam) in which case the estimated values of the Fisher coefficient are equal to one. The results for Malaysia, Singapore and Thailand in particular and those for South Africa and Egypt are consistent with those reported by Al-Khazali and Pyun (2004) and Alagidede and Panagiotidis (2010), respectively. For example, the results reported by Al-Khazali and Pyun (2004) show that the estimated values of the Fisher coefficient (θ) for Malaysia, Singapore and Thailand are 1.21, 1.28 and 1.07, respectively, whereas those reported by Alagidede and Panagiotidis (2010) are 2.26 and 0.22 for South Africa and Egypt, respectively.

The results based on the error-correction representations normalized on stock prices as well as goods prices, as reported in Table 4, show that the estimated speed of adjustment coefficients have the expected signs in all the cases, although they indicate that the speed of adjustment towards long-run equilibrium is very slow. The estimated values of the speed of adjustment coefficients in the error-correction model normalized on stock prices range from −0.047 (Vietnam) to −0.001 (Latvia), whereas those in the error-correction model normalized on goods prices range from 0.001 (Thailand) to 0.007 (Singapore). This implies that it takes quite a long time for stock prices (goods prices) to revert to their long-run equilibrium following an unexpected movement in goods prices (stock prices). The results of the error-correction representations are also consistent with those reported by Al-Khazali and Pyun (2004) and Alagidede and Panagiotidis (2010).

The results of the Pesaran et al. (2001) ARDL bounds test of cointegration, as reported in Table 5, are based on an unrestricted error-correction model specified as follows:

$$\Delta s_t = \alpha_0 + \alpha_1 s_{t-1} + \alpha_2 p_{t-1} + \sum_{i=1}^{p} \beta_i \Delta s_{t-i} + \sum_{i=0}^{q} \gamma_i \Delta p_{t-i} + \omega_t \tag{16}$$

where α_0 is the drift component, α_1 and α_2 are long-run multipliers, p and q are the lag lengths for Δs_{t-i} and Δp_{t-i}, respectively, selected by minimizing the Akaike information criterion (AIC), and ω_t is a white noise error term. If the model passes the diagnostic tests, we can then examine the long-run relationship between stock and goods prices by testing the null hypothesis of no cointegration, that is, $\alpha_1 = \alpha_2 = 0$, on the basis of the F-statistic. The calculated F-statistic is then compared with a pair of critical values (an upper and a lower bound) that are tabulated in Pesaran et al. (2001). Once the presence of cointegration is confirmed, the long-run relation can be established from the reduced form solution of equation (16), which boils down to equation (1), where $\gamma = -\alpha_0/\alpha_1$ and $\theta = -\alpha_2/\alpha_1$. Then, the error correction term, EC_t is constructed using the estimates of θ, and the corresponding restricted error correction model is given by

$$\Delta s_t = c + \sum_{i=1}^{p} \Psi_i \Delta s_{t-i} + \sum_{i=0}^{q} \lambda_i \Delta p_{t-i} + \pi EC_{t-1} + \omega_t \tag{17}$$

and tested to determine whether a valid error-correction representation exists.

The results based on Pesaran et al.'s (2001) test, as reported in Table 5, indicate that the null hypothesis of no cointegration can be rejected significantly for seven countries – China, Hong Kong, Hungry, Poland, Romania, Slovak Republic and Turkey – on the basis of the F-test. In four of the seven cases – China, Poland, Romania and Turkey – the null hypothesis of no cointegration can be rejected at the 5% significance level, whereas in the remaining three cases – Hong Kong, Hungary, and the Slovak Republic – it can be rejected at the 10% significance level.

Table 3 Cointegration tests based on Johansen's max eigenvalue and trace test statistics

Country	Max eigenvalue		Trace	
	None	At most one	None	At most one
Africa				
Ivory Coast	7.03	0.58	7.61	0.58
	[0.49]	[0.45]	[0.51]	[0.45]
Morocco	6.33	3.41[a]	9.75	3.41[a]
	[0.57]	[0.06]	[0.30]	[0.06]
Nigeria	11.07	0.47	11.54	0.47
	[0.15]	[0.49]	[0.18]	[0.49]
South Africa	17.65*	0.43	18.08*	0.43
	[0.01]	[0.51]	[0.02]	[0.51]
Tunisia	8.62	1.62	10.24	1.62
	[0.32]	[0.20]	[0.26]	[0.20]
South Asia				
Bangladesh	5.34	0.35	5.69	0.35
	[0.70]	[0.55]	[0.73]	[0.55]
India	7.59	2.73	10.32	2.73
	[0.42]	[0.10]	[0.26]	[0.10]
Pakistan	7.41	1.74	9.15	1.74
	[0.44]	[0.19]	[0.35]	[0.19]
Sri Lanka	6.92	2.24	9.16	2.24
	[0.50]	[0.13]	[0.35]	[0.13]
Southeast Asia				
Indonesia	8.50	1.16	9.66	1.16
	[0.33]	[0.28]	[0.31]	[0.28]
Malaysia	13.54[a]	0.30	13.84[a]	0.30
	[0.06]	[0.58]	[0.09]	[0.58]
Singapore	15.76*	0.00	15.76*	0.00
	[0.03]	[0.97]	[0.05]	[0.97]
Thailand	13.13[a]	2.96[a]	16.08*	2.96[a]
	[0.08]	[0.09]	[0.04]	[0.09]
Vietnam	13.47[a]	0.34	13.81[a]	0.34
	[0.07]	[0.56]	[0.09]	[0.56]
Middle East				
Bahrain	8.22	0.16	8.38	0.16
	[0.36]	[0.69]	[0.43]	[0.69]
Egypt	15.56*	0.04	15.60*	0.04
	[0.03]	[0.84]	[0.05]	[0.84]
Jordan	10.29	0.95	11.23	0.95
	[0.19]	[0.33]	[0.20]	[0.33]
Kuwait	12.19[a]	0.90	13.09	0.90
	[0.10]	[0.34]	[0.11]	[0.34]

The results of the long-run relationship and the ECM corresponding to the selected ARDL model for the 7 countries in which case stock and good prices are cointegrated are reported in Table 6.

These results show that the coefficients of long-run Fisher relation are correctly signed and statistically significant in all cases. The results of the ECM models are in line with expectations, since the coefficients of the lagged error-correction term, EC_{t-1} are negative and statistically significant across the board at the 5% significance level or better.

For the three South American countries (Brazil, Mexico and Peru) where both stock and goods prices are I(0) processes, we estimate equation (1) by fitting an autoregressive-moving average [ARMA(p, q)] model for the residuals to control for autocorrelation.[22] The results in Table 7 show that estimated values of the Fisher coefficient (θ) are correctly signed across the board; nonetheless, these values lack statistical significance, except for Brazil, where the Fisher coefficient is highly significant at any reasonable significance level.

Overall, the GFH performs better on the basis of the model specification using level data on stock and goods prices rather than on the basis of the model specification using first-difference data on the underlying variables. This is because the results are supportive of the long-run relationship in one-third of the cases, whereas they are supportive of the short-run relationship in one-tenth of the cases. This implies that common stocks provide a good hedge against inflation in the long rather than the short run and that there are deviations from the GFH in the short run.

5 Concluding remarks and discussion

This paper investigates the validity of the relationship between stock and goods prices, as embedded in the GFH, to ascertain whether investment in common stocks provides a hedge against inflation in emerging market countries. An attempt is made to present a coherent and somewhat chronological review of the vast literature to highlight the existing empirical evidence regarding the underlying relationship, especially for the stock markets of major developed and emerging market countries.

The bulk of the empirical work conducted in the 1970s, 1980s, and early 1990s on the stock markets of major developed market countries has produced evidence generally showing that nominal stock returns and inflation rates are inversely related, implying the invalidity of the GFH in the short run and failure of the proposition that investment in common stocks provides any hedge against inflation. Subsequently, a number of studies were carried out in an attempt to rehabilitate the GFH by looking into some neglected aspects of the empirical testing, such as the use of long-run data of a century or two and a long investment horizon, alternative cointegration tests, quantile regression and the index of the highest beta stocks. Overall, the findings from a considerable body of empirical work focusing on developed market countries are mixed, as they are highly sensitive to the model specifications used, the stock markets examined and the econometric procedures applied. These results can be summarized as follows. First, there is evidence supporting the rehabilitation of the GFH when examined over a century-long

22 Barnes et al. (1999, p. 747) “found that an ARMA (2, 1) process for the error term provided ‘the best fit.’”

Table 3 (continued)

Country	Max eigenvalue		Trace	
	None	At most one	None	At most one
Oman	14.12*	3.33[a]	17.44*	3.33[a]
	[0.05]	[0.07]	[0.03]	[0.07]
Qatar	7.00	1.64	8.65	1.64
	[0.49]	[0.20]	[0.40]	[0.20]
Saudi Arabia	6.87	0.09	6.96	0.09
	[0.50]	[0.76]	[0.58]	[0.76]
United Arab Emirates	11.37	1.40	12.78	1.40
	[0.14]	[0.24]	[0.12]	[0.24]
Eastern Europe				
Baltic states				
Latvia	22.41*	4.16*	26.57*	4.16*
	[0.00]	[0.04]	[0.00]	[0.04]
Lithuania	16.81*	0.87	17.68*	0.87
	[0.02]	[0.35]	[0.02]	[0.35]
Former Soviet states				
Kazakhstan	6.02	1.25	7.28	1.25
	[0.61]	[0.26]	[0.55]	[0.26]
Ukraine	3.24	0.16	3.41	0.16
	[0.93]	[0.68]	[0.95]	[0.68]
Central Europe				
Croatia	10.16	3.25[a]	13.41[a]	3.25[a]
	[0.20]	[0.07]	[0.10]	[0.07]
Southeastern Europe				
Bulgaria	12.31[a]	6.97*	19.28*	6.97*
	[0.10]	[0.01]	[0.01]	[0.01]

p-values in []. Max eigenvalue and Trace test *p*-values are from Mackinnon et al. (1999); lag order is 4 for all markets; unrestricted constant specification. * and [a] denote statistical significance at the 5 and 10% levels, respectively

(or two-century-long) period and a long investment horizon, since stock returns and inflation rates tend to be positively correlated (e.g., Boudoukh and Richardson 1993; Schotman and Schweitzer 2000; Lothian and McCarthy 2001). Second, the majority of the studies employing cointegration analysis (e.g., Anari and Kolari 2001; Al-Khazali and Pyun 2004; Luintel and Paudyal 2006) report findings indicating that stock prices tend to keep pace with goods prices in the long run and that the long-run elasticities of stock prices with respect to goods prices exceed unity. However, the results based on error-correction models indicate that the speed of adjustment towards the long-run relationship is quite slow, implying that stock prices take a very long time to return to their long-run relationship following unexpected movements in goods prices. Third, there is some evidence (e.g., Barnes et al. 1999; Ahmed and Cardinale 2005; Li et al. 2010) to suggest that the inflation-stock returns nexus varies across different inflationary

Table 4 Long-run relationship between stock prices and goods prices

Country	Long-run equation		Speed of adjustment		Restriction
	γ	θ	η_1	η_2	$\theta = 1$
Africa					
South Africa	−0.23	2.29*	−0.009	0.006*	16.66*
		(17.93)	(−0.46)	(4.18)	[0.00]
Southeast Asia					
Malaysia	−2.15	2.03*	−0.031*	0.002*	7.30*
		(7.44)	(−2.33)	(2.65)	[0.01]
Singapore	−0.81	1.88*	−0.044[a]	0.007*	3.99*
		(4.87)	(−1.84)	(3.65)	[0.05]
Thailand	−2.17	2.01*	−0.006	0.001*	3.34[a]
		(4.27)	(−0.86)	(3.54)	[0.07]
Vietnam	2.57	0.77*	−0.047*	0.004*	0.81
		(3.06)	(−2.41)	(2.75)	[0.37]
Middle East					
Egypt	5.79	0.51	−0.004	0.003*	
		(1.01)	(−0.46)	(3.81)	
Kuwait	5.31	0.71	−0.004	0.002*	
		(0.80)	(−0.68)	(3.36)	
Oman	−0.33	1.96*	−0.010	0.004*	2.40
		(4.53)	(−0.63)	(3.59)	[0.12]
Eastern Europe					
Baltic states					
Latvia	−2.71	1.94*	−0.001	0.006*	7.37*
		(6.33)	(−0.04)	(4.72)	[0.01]
Lithuania	−9.21	3.29*	−0.003	0.003*	7.53*
		(4.66)	(−0.23)	(4.08)	[0.01]

t-statistics are in (), *p*-values are in [], and the lag length is 4 for all markets. The restrictions imposed on the long-run equilibrium relation are tested using the likelihood ratio (LR) statistics with $\chi^2(1)$ degrees of freedom. * and [a] denote statistical significance at the 5 and 10% levels, respectively

regimes: a significantly positive relation between stock returns and inflation for high-inflation countries and an insignificant relation between stock returns and inflation in low-inflation countries. Third, some researchers (e.g., Bampinas and Panagiotidis (2016)) argue that the hedging ability of stocks can be enhanced by constructing investment portfolios of stocks with higher long-run betas as a part of an asset selection and allocation strategy. Finally, some studies (e.g., Gregoriou and Kontonikas 2010) employing alternative panel cointegration tests have produced results that lend strong support to the existence of cointegration between stock and goods prices, with Fisher coefficient estimates ranging between values less than and greater than one.

In contrast, the findings emerging from the massive amount of empirical work focusing on the stock markets of emerging market countries can be summarized as

Table 5 ARDL cointegration test results

Country	Selected model	F-statistic
East Asia		
Hong Kong	(2,0)	5.29[a]
Korea	(1,7)	4.48
Taiwan	(1,5)	3.92
China (Mainland)	(12,0)	9.35*
Southeast Asia		
Philippines	(2,0)	2.36
Middle East		
Israel	(5,4)	4.46
Turkey	(1,0)	8.17*
Eastern Europe		
Baltic states		
Estonia	(8,0)	3.82
Former Soviet states		
Russian Federation	(4,5)	2.98
Central Europe		
Czech Republic	(2,10)	4.07
Hungary	(2,1)	5.28[a]
Poland	(11,9)	8.59*
Slovak Republic	(6,1)	5.29[a]
Slovenia	(4,8)	3.46
Southeastern Europe		
Greece	(4,0)	0.79
Romania	(5,7)	6.07*
South America		
Chile	(1,0)	1.42
Colombia	(1, 0)	4.53

lag order of the $ARDL(p, q)$ model is determined by minimizing the Akaike information criterion (AIC); p and q are the lag length for Δs_{t-i} and Δp_{t-i}, respectively; unrestricted constant specification and no trend. The critical values are from Pesaran et al. (2001, p. 300), Table CI(iii): Case III: Unrestricted intercept and no trend with one regressor case. * and [a] denote statistical significance at the 5 and 10% levels, respectively

follows. First, there is some evidence to suggest (e.g., Barnes et al. 1999; Choudhry 2001) that common stocks in the emerging market countries with high inflation provide investors with a better effective hedge against inflation. One possible reason why high inflation emerging market countries behave differently than developed market countries with respect to the inflation-stock return relationship is that money rather than real activity is the dominant cause of inflation in these countries. Second, there is some evidence (e.g., Alagidede 2009) to suggest that the inflation-hedging potential of common stocks increases as the investment horizon increases. Finally, the results from cointegration analysis also provide some support for the validity of the GFH for some emerging market countries (e.g., Alagidede and Panagiotidis 2010).

This paper also offers some new empirical evidence on the long-run inflation-hedging property of common stocks for emerging market countries. Johansen's (1988) and Johansen and Juselius's (1990) maximum likelihood method of cointegration is applied to 28 countries for which the order of integration of stock and goods prices is the same, whereas Pesaran et al.'s (2001) ARDL bounds test of cointegration is applied to

Table 6 Cointegrating regression and the ECM

Country	Long-run form		ECM form	
	γ	θ	c	π
East Asia				
Hong Kong	−0.55	2.23*	−0.02*	−0.04*
	(−0.56)	(9.97)	(−2.36)	(−3.26)
China (mainland)	−0.71	1.82*	−0.06*	−0.09*
	(−0.42)	(4.81)	(−3.66)	(−4.33)
Middle East				
Turkey	6.31*	1.01*	0.46*	−0.07*
	(32.94)	(25.07)	(4.29)	(−4.05)
Central Europe				
Hungary	1.94*	1.74*	0.10*	−0.05*
	(2.38)	(9.30)	(3.27)	(−3.25)
Poland	−1.92	2.69*	−0.15*	−0.08*
	(−0.76)	(4.91)	(−4.13)	(−4.15)
Slovak Republic	−2.13	1.74*	−0.05*	−0.02*
	(−0.70)	(2.49)	(−2.94)	(−3.26)
Southeastern Europe				
Romania	4.43	0.96[a]	0.23*	−0.05*
	(1.61)	(1.65)	(3.56)	(−3.49)

t-statistics in (). * and [a] denote statistical significance at the 5 and 10% levels, respectively

18 countries for which the order of integration is not the same. Furthermore, regression analysis is applied to three countries – Brazil, Mexico and Peru – for which both the underlying series are stationary. The results based on Johansen's (1988) method of cointegration reject the null hypothesis of no cointegration between stock and goods

Table 7 ARMA conditional least squares regression

Coefficient/Country	Brazil	Mexico	Peru
γ	9.55*	10.54*	2.70
	(4.27)	(4.84)	(0.46)
θ	0.61*	0.23	1.39
	(7.28)	(0.58)	(1.12)
R^2	0.99	0.99	0.99
DW	2.04	1.98	2.01
B. G(2)	0.42	2.43	3.32
	[0.81]	[0.30]	[0.19]
RESET	2.26	8.01*	1.85
	[0.13]	[0.00]	[0.17]

DW = Dubin–Watson for first order residuals correlation and B.G = Breush–Godfrey test for higher-order residuals correlation. RESET = Ramsey's functional misspecification test. The residuals series for Brazil, Mexico and Peru are characterized by AR(1), ARMA(3,2) and ARMA(3,1) processes, respectively. *t*-statistics are reported in (), while *p*-values are reported in []. * denotes significance at the 5% level prices in only 10 cases, whereas those based on Pesaran et al.'s (2001) test of cointegration reject it in 7 cases. Overall, the results are consistent with the GFH in 10 countries for which cointegration is established; the null hypothesis of no cointegration is rejected at the 5% significance level for 6 countries, whereas the underlying null hypothesis is rejected at the 10% significance level for the remaining countries. The results of the coefficient restriction test based on the maximum likelihood ratio cannot reject the null hypothesis of one-to-one correspondence between stock and goods prices in two cases (Oman and Vietnam). The results imply that common stocks provide a superior hedge against inflation in six cases (South Africa, Singapore, Malaysia, Lithuania, Thailand and Latvia) and a perfect hedge in two cases (Oman and Vietnam). The results of the error-correction model suggest that stock prices take a very long time to return to their long-run equilibrium relationship with goods prices. Our results regarding the long-run relationship and error-correction representation are generally consistent with those of the previous studies conducted, inter alia, by Al-Khazali and Pyun (2004), Alagidede and Panagiotidis (2010) and Hassan et al. (2015). The Fisher coefficients are also consistent with those reported by Al-Khazali and Pyun (2004) and Alagidede and Panagiotidis (2010) for Malaysia (1.21), Singapore (1.28) and Thailand (1.07) and for South Africa (2.26) and Egypt (0.22), respectively. Overall, these findings lend some support to the GFH in the long run, implying that common stocks provide a hedge against inflation in 18 countries. One implication that emerges from these results is that monetary authorities are able to control inflation by reducing the nominal interest rates in the countries in which the GFH holds in the long run. Another implication is that monetary growth is the main driver of the long-run inflation rate in these countries.

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