

**THE LONG RUN RELATIONSHIP BETWEEN STOCK PRICES AND GOODS
PRICES: NEW EVIDENCE FROM PANEL COINTEGRATION**

Andros Gregoriou

Brunel Business School, Economics and Finance Section

Alexandros Kontonikas *

Department of Economics, University of Glasgow

* Corresponding author. Tel: +44 (0) 141 3306866; Fax. +44 (0) 141 3304940; E-Mail: a.kontonikas@lbss.gla.ac.uk

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Abstract

We examine the long run relationship between stock prices and goods prices to gauge whether stock market investment can hedge against inflation. Data from sixteen OECD countries over the period 1970-2006 are used. We account for different inflation regimes with the use of sub-sample regressions, whilst maintaining the power of tests in small sample sizes by combining time-series data across our sample countries in a panel unit root and panel cointegration econometric framework. The evidence supports a positive long-run relationship between goods prices and stock prices with the estimated goods price coefficient being in line with the generalized Fisher hypothesis.

JEL Classification: C33, G12

I. Introduction

Recent increases in energy and food prices along with the gradual evaporation of the inflation-calming effects of the 'globalization' supply shock that major economies have been enjoying ever since the early 1990s threaten to put global inflation in an upward trajectory. Although few would argue that a return to the highly inflationary 1970s is possible, given the major differences in policy regimes and underlying economic systems, nevertheless from an investor's point of view a re-examination of whether stock prices maintain their value relative to goods prices becomes increasingly important. According to the generalized Fisher hypothesis (GFE) since stocks represent claims to real assets their real rate of return should be uncorrelated to the underlying inflation rate, a prediction consistent with the classical view of mutually independent nominal and real sectors (Fisher 1930).

The perception of stocks as inflation-hedging investment was challenged by several empirical studies that offered compelling evidence supporting inflation's negative effect on short horizon (holding period of one year or less) stock returns. However, given that the Fisher hypothesis is an equilibrium relationship expected to hold in the long-run, the apparent failure to verify it using short-horizon regressions is not all surprising. Indeed, existing empirical evidence using the long-horizon regression methodology supports the existence of a positive long-run relationship between stock returns and inflation with estimated coefficients broadly in line with the GFE (see among others, Boudoukh and Richardson 1993). Nevertheless, since goods prices and stock prices are both known to be integrated processes with infinitely long memory, estimating regressions in terms of their first or higher order differences (corresponding to increasing holding horizons) implies that long-run information is only partially accounted for (Anari and Kolari 2001).

Hence, following developments in the econometric modeling of non-stationary time series recent literature on the long-run hedging properties of stock market investment has focused on modeling the levels of goods and stock prices using the cointegration framework developed by

Johansen (1988). Luintel and Paudyal's (2006) results indicate that UK stocks provide a good inflation-hedge over the long-run after allowing for structural breaks in the cointegrating relationship.

In this paper we undertake an alternative approach in tackling the possibility of structural change in the long-run relationship between stock prices and goods prices. Our approach entails conducting the empirical analysis not only across the full sample (1970-2006), but also across three sub-periods that correspond to three radically different inflation regimes in our dataset of sixteen OECD countries: high inflation (1970-1979), inflation moderation (1980-1989), inflation control (1990-2006). This will allow us to examine the process of disinflation affected the long-run elasticity of stock prices with respect to goods prices.¹

Breaking down the overall sample in three sub-samples significantly reduces the length of the dataset and it is well known that the power of time-series unit root and cointegration tests is conditional upon the use of long-span data (see among others, Zhou 2001).² Hence, unlike previous studies on the GFE, our unit root and cointegration analysis will be analyzed within a panel framework to utilize the dataset in the most efficient manner. We are the first study to our knowledge that examines the long-run relationship between stock prices and goods prices using panel cointegration. We will employ the panel unit root test established by Maddala and Wu (1999) and panel cointegration tests developed by Levin et al. (2002), Harris and Tzavalis (1999) and Maddala and Kim (1998). Cointegrating vectors are estimated using the fully modified OLS estimation technique for heterogeneous cointegrated panels developed by Pedroni (2000). This methodology allows consistent and efficient estimation of cointegrating vectors. In addition, it deals with the possible endogeneity in

¹ Blair Henry (2002) uses a dummy-variable regression methodology to estimate the impact of a disinflation policy announcement on stock returns. He finds that while the stock market significantly appreciates when countries attempt to control annual inflation rates in excess of 40 percent, there is a zero average stock market response when pre-stabilization inflation is less than 40 percent.

² Hakio and Rush (1991, p.572) point out that "cointegration is a long-run property and thus we often need long spans of data to properly test it".

the stock price – goods price relation and it encapsulates the time-series properties of the data in that integration-cointegration properties are explicitly accounted for.

Our dataset includes countries that have adopted inflation targeting monetary policy regimes at some point over the 1990s or early 2000's. An interesting question for stock market investors in terms of international investment allocation is whether the inflation-hedging property of stocks is affected by the underlying monetary policy regime. Therefore, in order to provide an answer to the aforementioned question we will break down the overall panel in two sub-panels: a panel including inflation targeting countries and a panel consisting of the remaining countries. The results from this analysis are also relevant to the discussion about the broader potential benefits of inflation targeting, given that if the long-run Fisherian elasticity of stocks is higher for targeting countries, this could be interpreted as an additional benefit from adopting targeting policies. Finally, we arrange the sample in a panel of high inflation countries and a panel of low inflation countries. This classification will enable us to further investigate whether the long-run relationship between stock prices and goods prices is affected by the underlying rate of inflation.³

II. A Review of the Generalized Fisher Hypothesis

The relationship between stock prices and goods prices has been the subject of extensive theoretical and empirical research over the last three decades. The GFE predicts a positive one-to-one ex ante relationship between stock returns and inflation making stocks a good hedge against inflation in the long-run.⁴ The main finding from early studies which focused on the US stock market was that,

³ There is some evidence indicating a positive response of stock market returns to inflation in high-inflation countries (see among others, Choudhry 2001). These findings, however, stem from short-horizon regressions and do not reveal information about the long-run hedging properties of stocks.

⁴ A review of contrarian arguments about the long-run effect on inflation on the stock market can be found in Barnes et al. (1999). They argue that high and sustained inflation has negative long-run consequences for the development of credit and

nominal and/or real short horizon stock returns were strongly negatively related to contemporaneous inflation, lagged inflation, and proxies of expected inflation, unexpected inflation, and the change in expected inflation (see among others, Fama and Schwert 1977). Further evidence suggesting a negative effect of inflation on share returns was provided in the context of multi-country studies (see among others, Barnes et al. 1999).

Various hypotheses have been proposed to explain the puzzling short-run findings involving e.g. inflation illusion (Modigliani and Cohn 1979). The debate on the empirical validity of these hypotheses is still active in the literature. Following attempts to provide an explanation for the puzzling negative short-run relationship between stock returns and inflation the literature has since then moved towards investigating the long-run hedging properties of stocks. In order to recover the long run information, two alternative methodologies have been adopted: regressions of long holding-period stock returns on inflation using long sample periods that span more than a century, and cointegration analysis of stock prices and consumer prices, pioneered by Boudoukh and Richardson (1993), and Ely and Robinson (1997), respectively.

Boudoukh and Richardson (1993) provide evidence in favour of the generalised Fisher effect using long time-series of annual stock returns in the U.S. (1802-1990) and in the U.K. (1820-1988) and an Instrumental Variable (IV) approach whereas expected stock returns and inflation are replaced by their ex-post counterparts and instruments are employed to proxy the ex-ante values.⁵ However, the long-horizon IV regression methodology has been subject to some criticism. One of its shortcomings is the use of overlapping observations on stock returns and inflation due to higher-order differencing which results in residual serial correlation and necessitates a correction of the estimated covariance matrix. As Engsted and Tanggaard (2002) point out, the standard procedures for such correction, such

equity markets. Some general equilibrium models also exhibit an inability of stocks to hedge against inflation (see among others, Marshall 1992).

⁵ Campbell and Vuolteenaho (2004) attribute this finding to the diminishing, in the long run, effect of inflation illusion.

as the Newey and West (1987) approach, are known to be unreliable when the degree of time overlap becomes large at long investment horizons.

The main advantage of the alternative modeling approach that has been employed in the GFE empirical literature, i.e. cointegration analysis, is that it allows the researcher to fully utilize the long run information. Ely and Robinson (1997) employ Johansen's (1988) framework and fail to find evidence supporting the long-run relationship between stock prices and goods prices for the majority of the sixteen countries that they considered over the period 1957-1992. Even in the few cases when a cointegrating vector exists, the relationship between stock prices and goods prices is not one-to-one. In sharp contrast to the findings of Ely and Robinson, Anari and Kolari's (2001) results indicate that stock prices and goods prices are cointegrated in six major economies over the period 1953-1998 with the long-run goods price coefficient significantly exceeding unity in four out of six cases.⁶ Finally, Luintel and Paudyal (2006) use UK aggregate and disaggregate data (1955-2002) and find that after adjusting for structural shifts, the long-run elasticity of stock prices with respect to goods prices exceeds unity in almost all cases, with disaggregate results exhibiting considerable heterogeneity.

III. Data Description

Data were collected from Datastream for sixteen OECD countries: Austria, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, Switzerland, United Kingdom, and United States. The sample period under investigation is January 1970 – June 2006, providing us with 438 monthly observations for goods prices, measured by the national consumer price index (P), and nominal stock prices, measured by the national stock price index (S).

⁶ The differences may be attributed to the different specifications that have been employed by the two studies. Particularly, Ely and Robinson (1997) add measures of real output and money in their empirical framework, arguing that these variables may affect the overall relationship between stock prices and goods prices. As Anari and Kolari (2001) point out, Ely and Robinson (1997) do not conduct bivariate tests of the long-run Fisher effect and do not report Fisher coefficients.

During the sample period, the growth rate of nominal stock prices exceeded that of goods prices in all sample countries, with the exception of Spain. Average monthly inflation ranges from 0.25% in Germany and Switzerland to 0.67% in Spain. Average monthly stock returns range from 0.42% in Germany to 1% in Finland. The average monthly growth rate of goods prices across our sample countries is 0.4%, while the corresponding nominal stock price growth rate is 0.7%, indicating a positive real stock price growth of 0.3%.

Our analysis will be conducted not only for the full sample 1970-2006, but also across three sub-periods: 1970-1979, 1980-1989, and 1990-2006. This will allow us to examine whether the full-sample results are robust to sample split corresponding to different inflationary environments. The relatively small number of time series observations in the three sub-samples will be dealt with by pooling the national datasets and utilizing panel approaches in our regression analysis. The chosen sub-periods correspond to three distinct inflation regimes. First, the highly inflationary 1970s over which the global economy experienced two oil shocks in 1973 and 1979. Second, the 1980s over which inflation started being moderated, and finally the last sub-period during which inflationary pressures were largely brought under control.⁷ Indeed, average monthly inflation rate across our sample countries has been progressively reduced from 0.7% (1970-1979), to 0.5% (1980-1989), and finally 0.2% (1990-2006). Our dataset includes seven countries that adopted inflation targeting at some point over the third sub-period: Canada, Finland, Norway, Spain, Sweden, Switzerland, and the United Kingdom.⁸ These countries are pooled in an unbalanced (due to the different dates of inflation targeting adoption) panel allowing us to examine whether response of stock prices to goods prices

⁷ Following a decline in all OECD countries inflation rates in the 1980s, a further decrease took place in the early-to-mid 1990s. Since then, inflation has exhibited relatively mild fluctuations around a low level.

⁸ Inflation targeting commenced in 1991.02, 2001.03, 1995.01, 2000.01, 1992.10, in Canada, Norway, Sweden, Switzerland, United Kingdom, respectively, and is currently still in place. Finland and Spain adopted inflation targeting over the 1990s prior to becoming members of the European Monetary Union, that is, over the periods, 1993.02-1998.12, and 1995.01-1998.12, respectively.

differs across the groups of inflation targeting versus non-targeting countries, respectively. Furthermore, we consider two additional panels: high inflation countries where annualized inflation rate during the full sample is greater than 6% (Ireland, United Kingdom, Italy, Spain), and low inflation countries with inflation less than 4% (Switzerland, Germany, Japan, Austria, and Netherlands).⁹ This will allow us to provide additional evidence on the robustness of the long-run relation between stock prices and goods prices.

IV. Econometric Framework and Results

In light of the GFE the international long run relationship between stock prices and goods prices can be expressed as follows:

$$s_{it} = \alpha_i + \beta_i p_{it} \tag{1}$$

where s_{it} , p_{it} denote the natural logarithm of stock prices and goods prices, respectively, for country i at time period t . Equation 1 implies a long-run relationship between stock prices and goods prices with causality running from the latter to the former. There is however some literature suggesting that stock price movements (along with developments in other asset prices such as house prices) affect output and inflation and that a broader price measure of prices that incorporates asset prices should be used by monetary policymakers (see among others, Goodhart 2001). These arguments imply that the long-run causal relationship between stock prices and goods prices points from the former to the latter. Thus, since the direction of long-run causality is generally not known *a priori*, in addition to Equation 1, we will consider the following potential equilibrium relationship:

$$p_{it} = \gamma_i + \delta_i s_{it} \tag{1.1}$$

The direction of the long-run causality will be determined on the basis of the results from cointegration analysis. The coefficient β in Equation 1 is the elasticity of stock prices with respect to

⁹ Alternative definitions of the high-low inflation countries, based e.g. on being above or below the median, produced similar results in the subsequent empirical analysis.

goods prices, i.e. the Fisher coefficient, indicating the percentage change in stock prices for every 1% change in goods prices. In order for common stocks to provide a long run hedge against inflation in a perfect market, β has to be at least equal to one¹⁰. Given that we are using stock prices and not total returns, we do not take into account the dividend component of stock returns. Ely and Robinson (1997) follow the same approach and point out that if stock price increases match goods price increases, it can be argued that stocks offer protection against inflation since the excluded dividends cannot be negative.

Luintel and Paudyal (2006) suggest regressing the first difference of log-stock prices (monthly nominal stock returns) on that of goods prices (monthly inflation) as a precursor to the cointegration analysis. The regression results (available upon request) indicate that in most sample countries (13/16) the inflation coefficient is statistically insignificant at the 5% level of significance. The inflation coefficient is negative in most cases (11/16), being positive and statistically different from zero only in Norway. The estimated value of the inflation coefficient in Norway is 0.5, significantly less than the value of unity that is implied by the GFE. Thus, overall, the results from first difference regressions do not lend support to the idea that stocks can hedge against inflation, which is in line with the findings from earlier short horizon studies. Moreover, our results do not support the idea that high inflation countries exhibit a positive stock returns-inflation relationship, since the estimated inflation regression coefficient in our four highest inflation countries is either statistically insignificant (Ireland, UK, Italy), or significantly negative at the 10% level (Spain).

¹⁰ As Luintel and Paudyal (2006) point out, the application of taxation on income from stocks would imply that β should be greater than one so that the long-run rate of return rate on common stocks exceeds the inflation rate at least by the tax rate. See Darby (1975) for the derivation of the tax-adjusted version of the Fisher hypothesis.

Panel unit root test results

In order to evaluate a possible long run relationship between stock prices and goods prices we need to first establish the order of integration of the variables. It is widely recognized that time series unit root tests may suffer from low power, especially with short spanned data (see among others, Pierse and Shell 1995). Hence we will consider a more powerful panel approach to examine the degree of non-stationarity in goods prices and stock prices. The notion that stock prices follow a random walk process and are therefore I(1) is generally taken as a stylized fact with existing empirical evidence overwhelmingly supporting it (see e.g. Anari and Kolari, 2001). However, there is lack of consensus in the empirical time-series literature on the order of integration of goods prices thereby motivating the panel approach in order to reduce the probability of spurious non-rejection of the null unit root hypothesis¹¹. We have employed a panel unit root test established by Maddala and Wu (1999), denoted as the MW statistic¹².

The MW statistic is given by $P = -2 \sum_{i=1}^N \ln p_i$ and combines the p-values from the individual ADF test statistics. The P test follows a χ^2 distribution with degrees of freedom twice the number of cross-section units, i.e. $2N$ under the null hypothesis of non stationarity. The MW panel unit root test results in Table 1 cannot reject the null-unit root hypothesis in levels, but do strongly reject it in first differences suggesting that both stock prices and goods prices are I(1) variables. The results are robust to the panel sample-split corresponding to distinct inflation regimes.

[INSERT TABLE 1 HERE]

¹¹ Moving away from time-series approaches and adopting more powerful panel unit root tests, Culver and Papell (1997) find that if national inflation rates are pooled the null hypothesis of panel unit root can be rejected.

¹² An alternative panel unit root test has developed by Im et al. (2003) which involves the averaging of individual ADF unit root test statistics. We have not computed this test for the following reasons. First, Breitung (1999) finds that the Im et al. (2003) test suffers from a dramatic loss of power when individual trends are included. Second, the MW test has the advantage over the Im et al. (2003) test that its value does not depend on different lag lengths in the individual ADF tests.

Panel cointegration tests results

Having established that stock prices and goods prices are I(1) variables, we proceed to examine whether there is long run co-movement amongst them. The Johansen (1988) time-series cointegration test can be severely distorted when the data-span is not long. Given that the data-span in our sub-samples is short¹³, we conduct three panel cointegration tests in order to circumvent the possibility of power problems in Johansen's time-series test. In particular, we employ the Levin et al. (2002), Harris and Tzavalis (1999), and Maddala and Kim's (1998) panel cointegration tests¹⁴. A limitation in both the Levin et al. and Harris and Tzavalis tests is that they do not allow for heterogeneity in the autoregressive coefficient of the panel. On the other hand, Maddala and Kim's Fisher's test does not assume homogeneity of coefficients in different countries because it aggregates the p -values of individual Johansen maximum likelihood cointegration test statistics. As Maddala and Kim on page 137 demonstrate if p_i denotes the p -value of the Johansen statistic for the i th unit, then it can be

shown that:
$$-2 \sum_{i=1}^N \log p_i \sim \chi^2_{2N}.$$

Panel cointegration tests results in Table 2 are very conclusive in identifying that stock prices and goods prices are cointegrated over the full sample period and the sub-samples, thereby supporting the robustness of existing time-series evidence. The reported results from Levin et al. and Harris and Tzavalis tests assume that the stock price is the dependent variable, while Maddala and Kim's test does not rely upon such an assumption. Replacing stock prices with goods prices as the dependent variable results into finding no-cointegration between the two variables (results not reported, available upon request). Thus, our findings suggest that a long-run relationship between stock prices and goods

¹³ Discussing the low power of time-series cointegration tests with short-span data, Hakkio and Rush (1991, p.572) argue that "testing a long-run property of the data with 120 monthly observations is no different than testing it with 10 annual observations".

¹⁴ For a detailed exposition of the Levin et al. (2002) and Harris and Tzavalis (1999) panel cointegration tests readers are referred to the appendix.

prices does exist with long-run causality pointing from the latter to the former, in line with the prediction of the GFE.

[INSERT TABLE 2 HERE]

Estimating the long run relationship between stock prices and goods prices

Given that stock prices and goods prices are cointegrated to fully evaluate the prediction of long-run hedging inherent in the GFE, we will estimate the long-run elasticity of stock prices to goods prices in Equation (1) using the Pedroni (2000) fully modified OLS heterogeneous cointegrated panel methodology. This methodology addresses the problem of non-stationary regressors, as well as possible endogeneity and simultaneous determination of stock and goods prices. The following system of equations is employed:

$$y_{it} = \alpha_i + x'_{it}\beta + u_{it} \quad (2)$$

$$x_{it} = x_{i,t-1} + e_{it} \quad (3)$$

where $\xi_{it} = [u_{it}, e'_{it}]$ is stationary with covariance matrix Ω_i . Pedroni (2000) applies a semi-parametric correction to the OLS estimator that eliminates second order bias caused by the endogeneity of the regressors in the panel. Essentially, he allows for heterogeneity in the short run dynamics and in the fixed effects of the panel. More specifically, the Pedroni (2000) estimator is:

$$\hat{\beta}_{FM} - \beta = \left(\sum_{i=1}^N \hat{\Omega}_{22i}^{-2} \sum_{t=1}^T (x_{it} - \bar{x}_t)^2 \right) \cdot \sum_{i=1}^N \hat{\Omega}_{11i}^{-1} \hat{\Omega}_{22i}^{-1} \left(\sum_{t=1}^T (x_{it} - \bar{x}_t) u_{it}^* - T \hat{\gamma}_i \right) \quad (4)$$

$$\hat{u}_{it}^* = u_{it} - \hat{\Omega}_{22i}^{-1} \hat{\Omega}_{21i}, \quad \hat{\gamma}_i = \hat{\Gamma}_{21i} + \hat{\Omega}_{21i}^0 - \hat{\Omega}_{22i}^{-1} \hat{\Omega}_{21i} \left(\hat{\Gamma}_{22i} + \hat{\Omega}_{22i}^0 \right) \quad (5)$$

where the covariance matrix can be decomposed as $\Omega_i = \Omega_i^0 + \Gamma_i + \Gamma_i$ where Ω_i^0 is the contemporaneous covariance matrix, and Γ_i is a weighted sum of autocovariances.

Table 3 reports fully modified OLS estimates of the long-run relationship between stock prices and goods prices for the panel as a whole over the full sample period and the three sub-samples. The reported Fisher coefficient is significantly different from zero at the 1% level in all cases suggesting that the stock market is affected by the developments in goods prices. The panel estimate of the long-run elasticity of stock prices with respect to goods prices is equal to 0.87 over the full sample period (1970-2006), which is lower in magnitude compared to the time-series based results in the literature where the reported coefficients generally exceed unity.

Focusing on the sub-sample regression results, it appears that the shift from a high inflation regime towards a low inflation regime was associated by an increase in the Fisher coefficient. Particularly, the goods price elasticity of stock prices exhibits its lowest value ($\beta=0.75$) during the highly inflationary 1970s but then increases over the 1980s ($\beta=0.88$), finally exceeding unity ($\beta=1.1$) during the last sub-period (1990-2006) of low and stable inflation. However, once the standard error of the estimates is taken into account within a *t*-test where the null hypothesis is that the Fisher coefficient is equal to one, the null cannot be rejected in all cases indicating support for the notion that common stocks provide a long run hedge against inflation. Our results do not support Darby's (1975) tax hypothesis since the estimated Fisher coefficients do not significantly exceed unity.

[INSERT TABLE 3 HERE]

In Table 4 we report fully modified OLS estimates of the long-run elasticity of stock prices to goods prices across targeting versus non-targeting countries and across high versus low inflation countries. Regarding the latter, the evidence points towards the same direction as the sub-sample evidence in Table 3, that is, lower inflation is associated with a Fisher coefficient of greater magnitude: $\beta=0.94$ in the panel with low-inflation countries and 0.69 in the high inflation group. Nevertheless, in both cases the estimated elasticity is statistically indistinguishable from unity thereby suggesting that stocks have served as inflation-hedging instrument in both high and low inflation countries. Finally, panel estimates of the Fisher coefficient indicate differences in its magnitude across

targeting versus non-targeting countries, with the group of targeters exhibiting a higher elasticity: $\beta=1.18$, as opposed to 0.99 for non-targeters. Once more, the null hypothesis of unity elasticity cannot be rejected thereby supporting the non-tax adjusted version of the GFE.¹⁵

[INSERT TABLE 4 HERE]

An important caveat of the empirical analysis is that all forms of cross-sectional dependency have been ignored. If there is cross-sectional dependence, then the panel cointegration tests depend on nuisance parameters associated with the cross-sectional correlation properties of the data, which means that the tests no longer have a limiting normal distribution. One possible solution is to derive critical values in the presence of non-normality by applying a wild bootstrap simulation (for more information on the wild bootstrap see among others, Arghyrou and Gregoriou 2007), which makes inference possible under cross-sectional dependence. For robustness we derived critical values from the wild bootstrap (results not reported, available upon request) and found that the bootstrap test performs well and the panel cointegration tests based on the normal distribution are robust to cross-sectional correlation.

V. Summary and Conclusions

In this paper we examine the long run relationship between stock prices and goods prices in order to determine whether stocks market investment can provide a hedge against inflation. We use data from sixteen OECD countries over the sample period 1970-2006. Preceding time-series based empirical evidence supports the existence of a positive cointegrating relationship between consumer prices and goods prices, however the previous literature has not explicitly accounted for the impact of changing

¹⁵ The non-rejection of the null hypothesis of unity Fisher coefficient in the cases of the all countries panel and inflation-targeting countries panel can be attributed to the relatively large standard errors associated with the Fisher coefficient. Boudoukh and Richardson (1993) using a different approach also report that their estimated coefficients were generally consistent with the Fisher model, albeit with relatively large standard errors.

inflation regimes. We encapsulate different inflation regimes with the use of sub-sample regressions, whilst maintaining the power of tests in small sample sizes by combining time-series data across our sample countries in a panel unit root and panel cointegration econometric framework. We find strong evidence in favor of a positive long-run relationship between goods prices and stock prices with long-run causality running from the former to the latter.

To estimate the long-run elasticity of stock prices with respect to goods prices, we employed the Fully Modified OLS panel estimator, which accounts for potential endogeneity, joint determination and non stationarity of the regressors. We found that the goods price elasticity of stocks increased in magnitude over time in line with the shift from a high to a low inflation regime. Further classifications of our sample countries reveal the estimated elasticity is greater in low inflation countries and inflation targeting countries. In all the cases that we considered, the null hypothesis of a one-for-one long-run relationship between stock prices and goods prices cannot be rejected.

Overall, our findings support the generalized Fisher hypothesis and are consistent with the view that stocks hedge against inflation in the long-run. Given the recent upward pressures in goods prices and the fear of a return to a state of higher inflation the empirical results of this paper are important for stock market investors and therefore cannot be ignored.

Appendix: Exposition of the Levin et al. and Harris and Tzavalis panel cointegration tests.

We compute the Levin et al. (2002) test by considering the following model:

$$y_{it} = \rho_i y_{i,t-1} + z'_{it} \gamma + u_{it} \quad (\text{A1})$$

Where z_{it} are deterministic variables, u_{it} is an error term with a mean of zero and a variance of σ^2 and

$\rho_i = \rho$. The test statistic is a t statistic on ρ given by

$$t_\rho = \frac{(\hat{\rho} - 1) \sqrt{\sum_{i=1}^N \sum_{t=1}^T \hat{y}_{i,t-1}}}{s_e} \quad (\text{A2})$$

Where $\hat{y}_{it} = y_{it} - \sum_{s=1}^T h(t,s) y_{is}$, $\hat{u}_{it} = u_{it} - \sum_{s=1}^T h(t,s) u_{is}$, $h(t,s) = z'_t \left(\sum_{t=1}^T z_t z'_t \right) z_s$,

$$s_e^2 = (NT)^{-1} \sum_{i=1}^N \sum_{t=1}^T \hat{u}_{it}^2,$$

and $\hat{\rho}$ is the OLS estimate of ρ . It can be shown that if there are only fixed effects in the model, then

$$\sqrt{NT}(\hat{\rho} - 1) + 3\sqrt{N} \rightarrow N\left(0, \frac{51}{5}\right)$$

and if there are fixed and time effects

$$\sqrt{N}(T(\hat{\rho} - 1) + 7.5) \rightarrow N\left(0, \frac{2895}{112}\right)$$

Second, we use the panel cointegration tests for equation (A1) by Harris and Tzavalis (1999). If there are only fixed effects in the model, then

$$\sqrt{N}\left(\hat{\rho} - 1 + \frac{3}{T+1}\right) \rightarrow N\left(0, \frac{3(17T^2 - 20T + 17)}{5(T-1)(T+1)^3}\right)$$

If there are fixed and time effects, then $\sqrt{N}\left(\hat{\rho} - 1 + \frac{15}{2(T+2)}\right) \rightarrow N\left(0, \frac{15(193T^2 - 728T + 1147)}{112(T+2)^3(T-2)}\right)$

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Table 1: Panel unit root test results

Time Period	Variables	MW Level	MW First Difference
1970-2006	Stock prices	15.23	56.78 ***
	Goods prices	17.84	62.23 ***
1970-1979	Stock prices	13.25	59.35 ***
	Goods prices	15.67	60.57 ***
1980-1989	Stock prices	17.65	56.25 ***
	Goods prices	18.32	59.86 ***
1990-2006	Stock prices	19.32	50.23 ***
	Goods prices	21.36	52.65 ***

Note: MW is the Maddala and Wu (1999) panel unit root test statistic. The critical values of the MW test are 37.57 and 31.41 at the 1% and 5% significant levels, respectively. *** signifies rejection of the null hypothesis of non stationarity at the 1% level of significance.

Table 2: Panel cointegration tests results

Test	Test Statistic			
	1970-2006	1970-1979	1980-1989	1990-2006
Levin-Lin-Chu (Fixed Effects Only)	-5.63***	-6.23***	-6.12***	-5.72***
Levin-Lin-Chu (Fixed and Time Effects)	-5.04***	-6.01***	-5.98***	-5.68***
Harris-Tzavalis ((Fixed Effects Only)	-9.32***	-9.37***	-9.21***	-9.94***
Harris-Tzavalis (Fixed and Time Effects)	-7.83***	-7.92***	-8.04***	-8.27***
Maddala-Kim (r=0)	59.55***	62.34***	60.23***	65.67***
Maddala-Kim (r≤1)	22.60	20.16	21.14	19.92

Note: The critical values of the Levin et al. (2002) and Harris and Tzavalis tests (1999) are -1.72, -1.78, respectively at the 5% level, and -2.40 and -2.72 at the 1% level, respectively. The critical values of Maddala and Kim's (1998) Fisher's χ^2 test are 37.57 and 31.41 at the 1% and 5% level respectively. Fisher's test is based on p-values from Johansen's likelihood cointegration methodology, therefore it applies regardless of the dependant variable. *** signifies rejection of the null hypothesis of no cointegration at the 1% level of significance, and r denotes the number of cointegrating vectors.

Table 3: Fully modified OLS estimates of the long-run relationship between stock prices and goods prices

Time Period	Estimate of β
1970-2006	0.87 (3.00) *** ^N
1970-1980	0.75 (3.45) *** ^N
1980-1990	0.88 (3.62) *** ^N
1990-2006	1.10 (3.84) *** ^N

Note: β is the long-run elasticity of stock prices with respect to goods prices estimated by Fully Modified OLS. Figures in parentheses are the values of the t -statistic associated with the null hypothesis that the β coefficient is equal to zero. *** signify rejection of the null hypothesis that the β coefficient is equal to zero at the 1% level of significance. ^N signifies non-rejection of the null hypothesis that the β coefficient is equal to one at the 10% level of significance.

Table 4: Fully modified OLS estimates of the long-run relationship between stock prices and goods prices in high vs. low inflation countries and inflation targeting vs. non-targeting countries

Time Period	Estimate of β		
	All countries	High inflation countries	Low inflation countries
1970-2006	0.87 (3.00) *** ^N	0.69 (2.86) *** ^N	0.94 (4.02) *** ^N
	All countries	Inflation targeting countries	Non-inflation targeting countries
1990-2006	1.10 (3.84) *** ^N	1.18 (3.14) *** ^N	0.99 (2.78) *** ^N

Note: See Table 3 notes. The all countries panel is a balanced panel including all sixteen sample OECD countries. The high (low) inflation countries panel is a balanced panel including countries where annualized inflation over the period 1970-2006 is greater (smaller) than 6% (4%). The inflation targeting countries panel is an unbalanced panel including countries that currently practice or have practiced inflation targeting.