

A Tale of Two Effects

by Paul Evans, Ohio State University,
and Xiaojun Wang, University of Hawaii-Manoa

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Abstract

This paper adopts a New Keynesian approach to analyze the relationship between nominal interest rates and prices. In this new framework, both a positive relation between interest rates and price *levels* (i.e., a positive Gibson effect) and a negative relation between interest rates and subsequent price *changes* (i.e., a negative Fama-Fisher effect) arise when money is supplied inelastically and prices are flexible. Such an economy is subject to Gibson's Paradox, a long-standing puzzle in monetary economics, and a novel paradox identified here, a Fama-Fisher Paradox. By contrast, economies characterized by elastic money and sticky prices are not so paradoxical since nominal interest rates are positively related to subsequent inflation and ambiguously related to the price level. Empirical analysis of nearly two centuries of data for ten countries supports the new theory.

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Corresponding Author: Paul Evans, Department of Economics, Ohio State University, Columbus, OH 43210. Phone: (614) 292-0072. <evans.21@osu.edu>

Xiaojun Wang, Department of Economics, University of Hawaii-Manoa, 2424 Maile Way, Saunders Hall Room 542, Honolulu, Hawaii, U.S.A. 96822. Phone: (808) 956-7721 <xiaojun@hawaii.edu>

1 Introduction

Empirical research on the relationship between the nominal interest rate and the price level has left monetary economists with two puzzles. On the one hand, in one of the most influential works on monetary economics, Irving Fisher (1930) suggested a positive relationship between the nominal interest rate and the growth rate of the price level. However, little empirical evidence has been found for such a relationship prior to World War II.¹ On the other hand, the nominal interest rate and price level are positively related before World War I and the relationship apparently disappeared after that. Because the prewar relationship seemed to be inconsistent with the neutrality of money, John Maynard Keynes (1930) dubbed it *Gibson's Paradox* and considered it to be “one of the most completely established empirical facts in the whole field of quantitative economics.”²

World War I marked not only the apparent disappearance of Gibson's Paradox but also the end of the classical gold standard. Building on these apparent facts, Robert Barsky and Lawrence Summers (1988) formulated a model to explain Gibson's Paradox as a phenomenon that depended on the peculiarities of the classical gold standard. They showed that the modest real interest shocks documented by Robert Shiller and Jeremy Siegel (1977) might have appreciably affected the demand for gold, inducing an appreciable change in the price level. If the shocks were also persistent, the inflation premium in nominal interest rates would be little affected. As a result, a positive real interest shock would increase both the nominal interest rate and the price level. With the demise of the classical gold standard, these mechanisms would no longer operate and therefore Gibson's Paradox would vanish.³

¹ See Richard Roll (1972) and Thomas Sargent (1973) for surveys of the evidence.

² Keynes (1930, v. 2, p. 198) believed that A. H. Gibson, a British financial journalist, had discovered the relationship first (Gibson, 1923), but Knut Wicksell had documented the same relation 16 years earlier (Wicksell, 1907).

³ Many prominent economists have also provided their own explanations of the Gibson's Paradox, including Fisher, Wicksell, Keynes, and Sargent, to name just a few. For details, see Milton Friedman and Anna Schwartz (1982). Gerald Dwyer (1984) documented the instability of Gibson's Paradox across different monetary systems in four countries.

This paper provides answers to the following three questions. First, what does theory predict the relationship to be between the nominal interest rate and the growth rate of the price level under a commodity standard? That is, are the “non-Fisherian” empirical findings for the period before World War II really due to poor data as alleged by Eugene Fama (1975)? Second, is the Gibson paradox only a gold-standard phenomenon? And finally, is it possible for the nominal interest rate to be related to the growth rate of the price level and also to the price level itself, and how are these two relationships connected?

We propose a New Keynesian framework that provides answers to these questions. In particular, our model is flexible enough to incorporate both the gold standard and fiat money standards as special cases. The model shows that the relationship between the nominal interest rate and the subsequent growth rate of the price level (henceforth, the Fama-Fisher effect) should be *negative* and the relationship between the nominal interest rate and the price level (henceforth, the Gibson effect) should be positive when money is inelastically supplied and prices are flexible. Because these conditions were met before World War I, a twin paradox (i.e. a negative Fama-Fisher effect *and* a positive Gibson effect) arose during this period.

Our analysis further suggests that the gold standard had relatively little to do with generating either Gibson’s Paradox or the Fama-Fisher Paradox and also that they might not actually have vanished after World War I. Data from the interwar period verify this conjecture: both Gibson’s Paradox and the Fama-Fisher Paradox characterize this period as well as the period before World War I.

A positive Fama-Fisher effect has emerged in the postwar period. In our model, this phenomenon reflects more elastically supplied money and stickier prices. For example, a positive Fama-Fisher effect arises if the central bank pursues a Taylor-type rule for the nominal interest rate and faces sufficiently sticky prices. Under these conditions, our model implies a Gibson effect of indeterminate sign. Empirically, its point estimates

continue to be mostly positive albeit statistically insignificant. In this sense, then, the postwar period is no longer paradoxical.

In section 2 we lay out a flexible model that allows the elasticity of the money supply and the flexibility of prices to vary. During the classical gold standard period, money was inelastically supplied and prices were flexible, while during the postwar period when fiat money standards prevailed, money was elastic and prices were sticky. The theoretical implications for the Fama-Fisher and Gibson effects are derived under each set of assumptions. Empirical evidence using pre-WWI data from seven industrialized countries is presented in section 3, followed by interwar and postwar evidence in section 4. Some final conclusions are drawn in section 5.

2 A Simple Model

We follow Richard Clarida, Jordi Galí and Mark Gertler (1999) in specifying the first two equations of the model as follows:⁴

$$i_t = E_t \Delta p_{t+1} + \varphi^{-1} E_t \Delta y_{t+1} + u_t, \quad \varphi > 0, \quad (1)$$

$$\Delta p_t = \beta E_t \Delta p_{t+1} + \kappa (y_t - y_t^*), \quad 0 < \beta < 1, \quad \kappa > 0; \quad (2)$$

In equations (1) and (2), t is a discrete index of time that assumes the values 1, 2, 3, ...; i_t is the nominal interest rate on a riskless one-period bond bought in period t and maturing in period $t+1$; p_t is the logarithm of the price level; y_t is the logarithm of output; y_t^* is the logarithm of the natural level of output; φ , β and κ are parameters; u_t is an error term that follows the first-order autoregressive process

$$u_t = \rho u_{t-1} + v_t, \quad |\rho| < 1; \quad (3)$$

v_t is an independently and identically distributed error term with a zero mean and finite variance; ρ is a parameter; and u_0 and p_0 are fixed initial values.

⁴ To avoid notational clutter, we assume that the intercepts in all equations of our model are zero.

Equation (1), the *New Keynesian IS Curve*, arises from the optimizing decisions of households and firms about when to consume and when to produce as mediated by the expected real interest rate *à la* Fisher. The error term u_t is a persistent shock to aggregate demand, which could arise from predictable changes in time preference, the share of output consumed by the government or adjustment costs for the capital stock.⁵

Equation (2), the *New Keynesian Phillips Curve*, arises from the optimizing decisions of firms in setting their sticky prices. The parameter κ is larger, the more flexible their prices are. In the limiting case of perfect flexibility ($\kappa \rightarrow \infty$), output always equals its natural level ($y_t = y_t^*$). Without loss of generality, we assume that firms experience no exogenous shocks to their markups.

For simplicity, we posit that the logarithm of the economy's natural level of output is a random walk:

$$\Delta y_t^* = v_t, \quad (4)$$

where v_t is an independently and identically distributed error term with a zero mean and finite variance and y_0^* is a fixed initial value. The error term v_t is a supply shock arising from exogenous changes in factor supplies and total factor productivity.

Households and firms also demand money. We assume that their demand for money takes the form

$$m_t^d = p_t + \eta y_t - \lambda i_t, \quad \eta, \lambda > 0, \quad (5)$$

where m_t^d is the logarithm of the stock of money demanded and η and λ are parameters.

For simplicity, we specify the elasticity of money demand with respect to the price level to be one even though it probably differed from one under the gold standard.⁶ We also

⁵ This is one interpretation of the shock considered by Robert Barsky and Lawrence Summers (1988).

⁶ Under the gold standard, money consisted of gold coins and claims on gold held in the vaults of private banks and governments. Gold also served nonmonetary functions, reflecting its shininess, malleability, ductility, and resistance to corrosion. Its price-level elasticity should be positive, however, since its relative

specify the income elasticity η and the interest semielasticity λ to be positive and constant. Without loss of generality, we omit any error term from equation (5) since its effect can be absorbed by μ_t in equation (6) below.

We assume that money is supplied according to the equation

$$\Delta m_t = \gamma(i_t - \tau E_t \Delta p_{t+1} - e_t) + \mu_t, \quad \gamma \geq 0, \quad \tau > 0, \quad (6)$$

where γ and τ are parameters, μ_t and e_t are independently and identically distributed error terms with zero means and finite variances, and m_0 is a fixed initial value. We specify the parameter γ to have been zero in the period before World War I and to have been positive and substantial in more recent periods.

The monetary arrangements in the United States before the founding of the Federal Reserve System were often criticized for failing to provide an “elastic currency.” In a sense, then, equation (6) with $\gamma = 0$ may be thought of as a perfectly inelastic currency, while an appreciable γ would represent an elastic currency. In the limit as γ approaches infinity, the currency becomes perfectly elastic.

On its face, our formulation of how the stock of money evolved under the commodity standards that prevailed before World War I would seem to be inconsistent with the fact that monetary metals were produced by private profit-maximizing firms. The production of these firms should have been decreasing in the price level since the price of their product was fixed and their variable costs were proportional to the price level. This reasoning suggests that Δm_t should have taken the form $\Delta m_t = -\delta p_t + \mu_t$ with $\delta > 0$ under the commodity standards in place before World War I. The model would, however, imply that p_t is covariance stationary, an implication that is not supported by the data; see the next section. Further evidence for the exogeneity of Δm_t during this period has been

price falls as the price level rises. Qualitatively, it is unimportant whether the elasticity is one or some other positive value.

provided by Barsky and Summers (1991), who document that technological shocks in the gold-mining industry dominated gold production.

Under fiat standards, we assume that central banks resist changes in the nominal interest rate and expected inflation. In particular, they meet a higher nominal interest rate or a lower expectation of inflation with a higher growth rate of the money supply; and conversely. In the extreme case in which γ approaches infinity, the central bank conducts its monetary policy using the forward-looking Taylor rule

$$i_t = \tau E_t \Delta p_{t+1} + e_t. \quad (7)$$

Clarida, Galí and Gertler (1998, 2000) and Christina Romer and David Romer (2002) have shown that forward-looking Taylor rules approximate the conduct of monetary policy during at least some periods and for many countries. Furthermore, many, if not most, central banks have long conducted monetary policy by manipulating the level of short-term interest rates. The error term e_t is a shock to monetary policy, which might arise if the central bank's target for inflation varies over time.

Equilibrium for the economy consists of a sequence of log price levels $\{p_t\}$, nominal interest rates $\{i_t\}$, log outputs $\{y_t\}$, and log stocks of money demanded and supplied $\{m_t^d, m_t\}$ in periods $t = 1, 2, 3, \dots$ such that equations (1)-(6) hold; the market for money is in equilibrium

$$m_t^d = m_t; \quad (8)$$

and u_0, p_0, y_t^* and m_0 take on fixed initial values.⁷

We now examine a few implications of the model in two special cases.

⁷ In the background, factors of production and bonds are also being demanded and supplied, and their demands and supplies are being equated.

2.1 Case I: $\gamma = 0$ and $\kappa \rightarrow \infty$

The period before World War I differs from more recent periods in two stark ways. First and most obviously, the gold-mining and to a lesser extent the silver-mining industries rather than central banks determined how money supplies evolved over time. Second, prices were highly flexible; see Charles Calomiris and Glenn Hubbard (1989) and Christopher Hanes and John James (2003) for evidence. We model these features of the period by equating γ to zero and letting κ approach infinity so that equation (2) reduces to $y_t = y_t^*$ and equation (4) becomes

$$\Delta y_t = v_t. \quad (4')$$

Equations (5), (6), (8), and (4') imply that

$$\Delta p_t = \lambda \Delta i_t + \mu_t - \eta v_t. \quad (9)$$

Substituting equations (9) and (4') into equation (1) then yields

$$i_t = \lambda E_t \Delta i_{t+1} + u_t$$

since $E_t \mu_{t+1} = E_t v_{t+1} = 0$. Solving this equation forward and using equation (3) gives us

$$i_t = \left(\frac{1}{1+\lambda}\right) \sum_{j=0}^{\infty} \left(\frac{\lambda}{1+\lambda}\right)^j E_t u_{t+j} = \left(\frac{1}{1+\lambda}\right) \sum_{j=0}^{\infty} \left(\frac{\lambda}{1+\lambda}\right)^j \rho^j u_t$$

or

$$i_t = [1 + \lambda(1 - \rho)]^{-1} u_t. \quad (10)$$

Finally, using equation (10) to eliminate Δi_t from equation (9), we have

$$\Delta p_t = [1 + \lambda(1 - \rho)]^{-1} \lambda \Delta u_t + \mu_t - \eta v_t. \quad (11)$$

It is straightforward to verify that both the Fama-Fisher Paradox and the Gibson Paradox arise in Case I. Consider the estimator obtained by applying OLS to the regression

$$\Delta p_t = \beta_F i_{t-1} + \varepsilon_{Ft} \quad (12)$$

where β_F is a parameter, ε_{Ft} is an error term, and F refers to Fama (1975), who first employed such regressions to test Fisher's *Theory of Interest* (1930).^{8,9} Equations (10) and (11) imply that the OLS estimator of β_F converges in probability to

$$\begin{aligned}\frac{\text{cov}(\Delta p_t, i_{t-1})}{\text{var } i_{t-1}} &= \frac{[1 + \lambda(1 - \rho)]^{-2} \lambda \text{cov}(\Delta u_t, u_{t-1})}{[1 + \lambda(1 - \rho)]^{-2} \text{var } u_{t-1}} \\ &= \frac{\lambda[\text{cov}(u_t, u_{t-1}) - \text{var } u_{t-1}]}{\text{var } u_{t-1}} = \lambda(\rho - 1)\end{aligned}$$

which is negative and certainly not one as Fama and Fisher hypothesized.

In contrast, consider the estimator obtained by applying OLS to the regression¹⁰

$$\Delta p_t = \beta_G \Delta i_t + \varepsilon_{Gt}, \quad (13)$$

where β_G is a parameter, ε_{Gt} is an error term, and G refers to Gibson (1923). The resulting estimator converges in probability to

$$\frac{\text{cov}(\Delta p_t, \Delta i_t)}{\text{var}(\Delta i_t)} = \frac{[1 + \lambda(1 - \rho)]^{-2} \lambda \text{var}(\Delta u_t)}{[1 + \lambda(1 - \rho)]^{-2} \text{var}(\Delta u_t)} = \lambda,$$

which is positive and equal to the interest semielasticity of money demand.

Consider a positive demand shock in period 1. The nominal interest rate rises, and given the inelastic money supply and flexible prices, the price level also rises. This causes a

⁸ Fama maintained that the expected real interest rate is always constant. Given this maintained assumption, we then have

$$\Delta p_t = \text{constant intercept} + i_{t-1} - (\Delta p_t - E_{t-1} \Delta p_t)$$

Under the assumption that expectations are rational, i_{t-1} is orthogonal to $(\Delta p_t - E_{t-1} \Delta p_t)$ so that applying OLS to equation (12) *cum* intercept should yield an estimator with a probability limit of one. Of course, in the model entertained here, the expected real interest rate is u_{t-1} , which is far from constant.

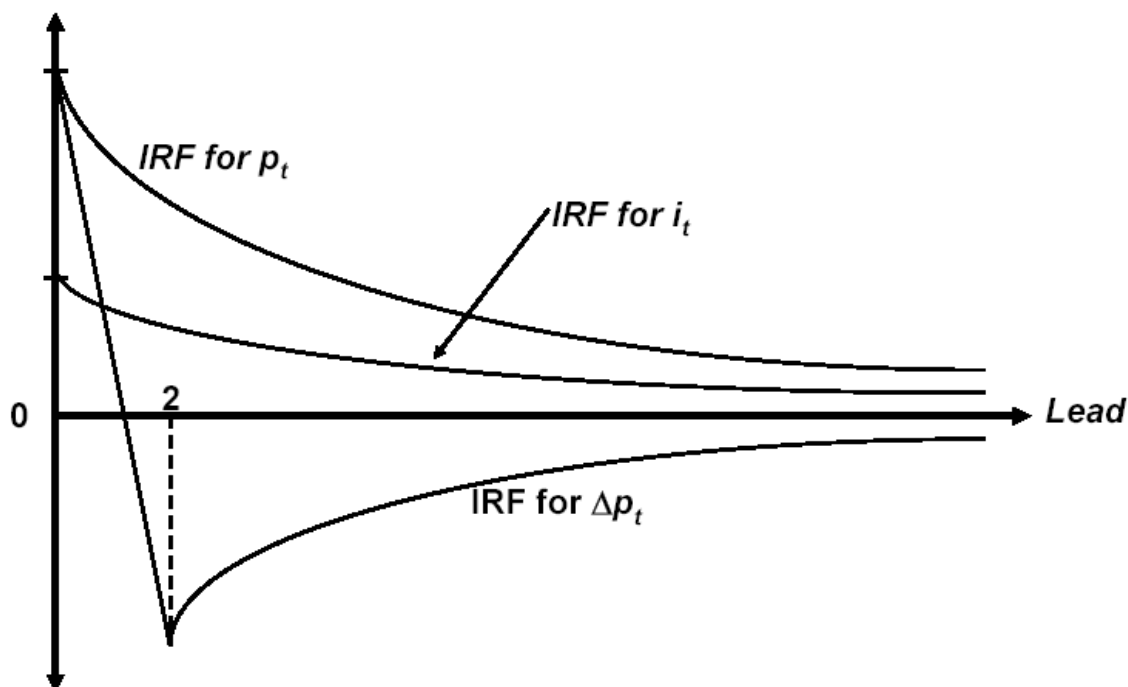
⁹ Fisher's identity, that the nominal interest rate equals the expected real interest rate plus the expected inflation rate, cannot be tested; it merely defines the expected real interest rate. Rather, Fama tested his and Fisher's maintained assumption that the expected real interest rate is constant while also maintaining the assumption of rational expectations.

¹⁰ Note that in the model considered here, the statistical properties of the estimator obtained by applying OLS to the regression

$$p_t = \beta_G i_t + \varepsilon_{Gt}$$

are nonstandard. The reason is that p_t is difference-stationary while i_t is covariance-stationary. To obtain an estimator with standard statistical properties, one must estimate a regression on differenced data; e.g., the one in the text.

Figure 1. Impulse Response Functions under Case I



contemporaneous positive correlation between the nominal interest rate and the price level; i.e., a positive Gibson's effect. However, because the demand shock is only temporary, the price level must subsequently return to its original level in the absence of other shocks. As a result, the inflation rate in subsequent periods must be less than average, resulting in a negative correlation between the current nominal interest rate and subsequent inflation, i.e. a negative Fama-Fisher effect. These responses are illustrated in Figure 1, which plots the impulse response functions (IRFs) of the nominal interest rate, price level and inflation rate.

As a matter of fact, the gold standard is not a necessary condition for a positive Gibson's effect and a negative Fama-Fisher effect. As long as money is inelastically supplied and prices are flexible, such a configuration of the two effects may arise in other periods. In this sense our approach is more general than that of Barsky and Summers (1988). Moreover, in the example given above, a positive Gibson's effect directly implies a

negative Fama-Fisher effect. This might explain why most work before Fama's (1975) failed to find a Fisherian coefficient even close to positive one.¹¹

2.2 Case II: $\gamma \rightarrow \infty$ and $\kappa < \infty$

We are modeling central banks as having followed Taylor rules of the form (7) in recent years. In addition, we are also assuming that prices are sticky to an appreciable extent. Letting $x_t \equiv y_t - y_t^*$, we can solve equation (2) forward to express the current inflation rate as follows:

$$\Delta p_t = \kappa \sum_{j=0}^{\infty} \beta^j E_t x_{t+j}. \quad (14)$$

Substituting equations (7) and (14) into equation (1), using the identity $x_t \equiv y_t - y_t^*$ and rearranging the resulting equation yields

$$x_t = E_t x_{t+1} + \varphi(1-\tau)\kappa \sum_{j=0}^{\infty} \beta^j E_t x_{t+1+j} + \varphi(u_t - e_t). \quad (15)$$

since $E_t \Delta y_{t+1}^* = 0$. The minimum-state-variable solution to this expectational difference equation takes the form¹²

$$x_t = [(1-\rho)(1-\beta\rho) + \rho\varphi(\tau-1)\kappa]^{-1} (1-\beta\rho)\varphi u_t - \varphi e_t. \quad (16)$$

Substituting equation (16) into equation (14) and using equation (3) gives us

$$\Delta p_t = \kappa [(1-\rho)(1-\beta\rho) + \rho\varphi(\tau-1)\kappa]^{-1} (1-\beta\rho)\varphi \sum_{j=0}^{\infty} \beta^j E_t u_{t+j} - \kappa\varphi \sum_{j=0}^{\infty} \beta^j E_t e_{t+j}$$

or

$$\Delta p_t = [(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1)]^{-1} \varphi u_t - \kappa\varphi e_t \quad (17)$$

since e_t is serially independent.¹³ Finally, substituting equation (17) into equation (7), we have

¹¹ The empirical work before Fama attempted to estimate this coefficient by regressing an interest rate on current and lagged inflation rates and summing the estimated coefficients. However, this coefficient sum has no straightforward interpretation if agents form rational expectations; see Sargent (1973) for further discussion of this point.

¹² One can verify this by substituting equation (16) into equation (15). If $\tau \leq 1$, there are multiple solutions to equation (15) including ones with sunspots. We do not consider those solutions in this paper.

$$i_t = \tau \left\{ \left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-1} \varphi E_t u_{t+1} - \kappa\varphi E_t e_{t+1} \right\} + e_t$$

or

$$i_t = \left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-1} \varphi\tau\rho u_t + e_t. \quad (18)$$

It is straightforward to verify that the Fama-Fisher effect is now positive and that the Gibson effect is of indeterminate sign. Equations (17) and (18) imply that the OLS estimator of β_F converges in probability to

$$\begin{aligned} \frac{\text{cov}(\Delta p_{t+1}, i_t)}{\text{var } i_t} &= \frac{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau \rho \text{cov}(u_{t+1}, u_t)}{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau^2 \rho^2 \text{var } u_t + \text{var } e_t} \\ &= \tau^{-1} / \left\{ 1 + \varphi^{-2} \tau^{-2} \rho^{-2} \left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^2 [\text{var } e_t / \text{var } u_t] \right\}, \end{aligned}$$

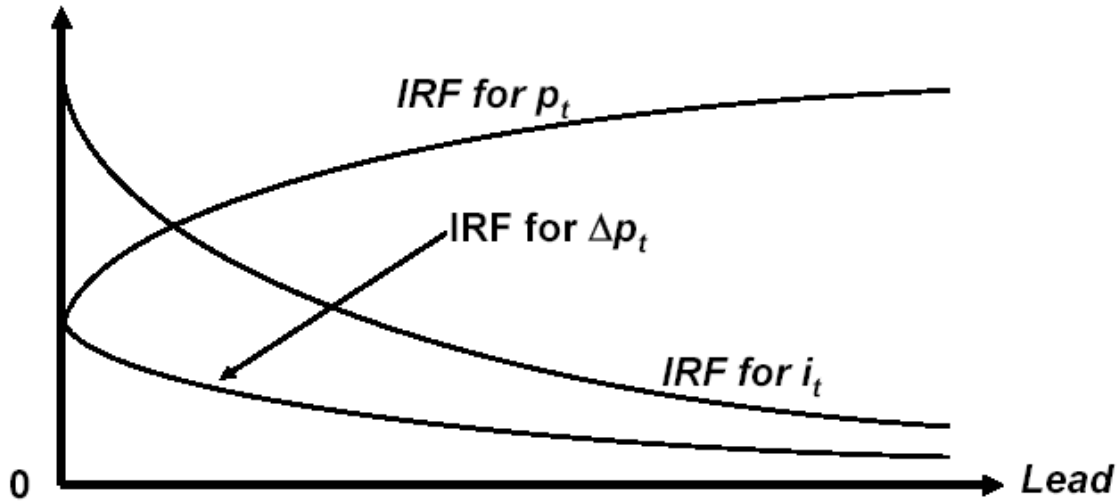
which lies strictly between zero and τ^{-1} and *a fortiori* between zero and one if $\tau > 1$. It is appreciable unless the central bank is highly aggressive ($\tau \gg 1$) or highly erratic ($\text{var } e_t \gg \text{var } u_t$). We therefore expect estimates of β_F to be positive, less than one, and highly significant under inconvertible fiat standards. In section 4, we present evidence consistent with this prediction of the theory.

In contrast, the OLS estimator of β_G converges in probability to

$$\begin{aligned} \frac{\text{cov}(\Delta p_t, \Delta i_t)}{\text{var}(\Delta i_t)} &= \frac{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau \rho \text{cov}(u_t, \Delta u_t) - \kappa\varphi \text{cov}(e_t, \Delta e_t)}{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau^2 \rho^2 \text{var}(\Delta u_t) + \text{var}(\Delta e_t)} \\ &= \frac{1}{2} \left\{ \frac{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau \rho [\text{var } u_t - \text{cov}(u_t, u_{t-1})] - \kappa\varphi \text{var } e_t}{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau^2 \rho^2 [\text{var } u_t - \text{cov}(u_t, u_{t-1})] + \text{var } e_t} \right\} \\ &= \frac{1}{2} \left\{ \frac{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau \rho (1-\rho) - \kappa\varphi (\text{var } e_t / \text{var } u_t)}{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau^2 \rho^2 (1-\rho) + (\text{var } e_t / \text{var } u_t)} \right\} \end{aligned}$$

¹³ Both inflation and interest rate have become much more persistent during the postwar period. Introducing persistency into e_t does not alter the main result but does complicate the analysis. We defer it to the appendix.

Figure 2. Impulse Response Functions for a Demand Shock under Case II

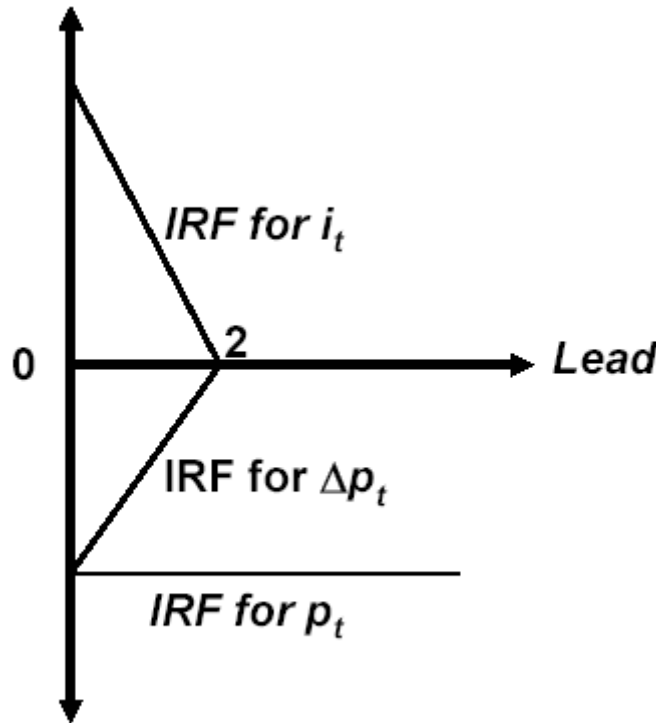


which can be negligibly positive or even negative if monetary policy is sufficiently variable relative to demand shocks; i.e., $\text{var } e_t / \text{var } u_t$ is sufficiently large.

Once again consider a positive demand shock in period 1. Because it is persistent and the central bank does not entirely offset its effects, the current inflation rate and next period's inflation rate rise. In anticipation of increased future inflation, the central bank raises the current nominal interest rate, leading to a positive correlation between the current nominal interest rate and next period's inflation rate; i.e., a positive Fama-Fisher effect. The increase in the inflation rate also results in an increase in the current price level, thereby contributing positively to the Gibson effect. These responses are illustrated in Figure 2.

This contribution is at least partly offset, and could be more than offset, by a second effect. Consider a positive shock to the nominal interest rate generated by a reduction in the central bank's desired inflation rate. This shock increases the current nominal interest

Figure 3. Impulse Response Functions for a Disinflationary Monetary-Policy Shock under Case II



rate and reduces the current inflation rate and price level, thereby contributing negatively to the Gibson effect. By contrast, because this shock is completely transitory, it does not contribute to the correlation between the current nominal interest rate and next period's inflation rate. These responses are illustrated in Figure 3.

3 Empirical Analysis for the Period before World War I

3.1 The Data

We estimate equations (12) and (13) using annual data on the wholesale price indices and short-term nominal interest rates of the United Kingdom, the United States, France, Belgium, Germany, the Netherlands and Switzerland for periods when those countries

adhered to commodity standards and were not in political turmoil.¹⁴ For the United Kingdom, France, Belgium, and Germany the wholesale price indices comes from Table I1 of Brian Mitchell (1992). The short-term interest rates for the United Kingdom, France, Belgium, Germany, Holland, and Switzerland come from Tables 23, 27, 29, 31, 33, 34, 61, 63, 65, 67, 69, and 71 of Sidney Homer and Richard Sylla (1991). For the United States, the wholesale price index comes from Table 1 of George Warren and Frank Pearson (1935) and the short-term interest rate comes from Table 10 of Frederick Macaulay (1938). The US data were originally monthly; we averaged them to obtain our annual series.

Wholesale price indices are not available for the Netherlands and Switzerland. We proxy the Dutch wholesale price index with a geometric average of those for Belgium, France, Germany and the United Kingdom and the Swiss wholesale price index with a geometric average of those for France and Germany. These countries are nearby and are assumed to have experienced similar price-level movements.

By far our longest sample is for the United Kingdom, which adhered to the gold standard from shortly after passage of the Resumption Act of 1819 until World War I began in August 1914. The years 1825-1913 comprise our sample since the open-market rate of discount is not available before 1824.

The sample periods for the other countries are shorter: Belgium, 1849-1913; France, 1873-1913; Germany, 1872-1913; the Netherlands, 1874-1912; Switzerland, 1851-1913; and the United States, 1879-1913. These sample periods are dictated partly by the availability of data but primarily by the narratives provided by Homer and Sylla, who detail periods of political turmoil in these countries.

¹⁴ Many of these countries were on bimetallic (gold-silver) standards before the 1870s. For the purposes of this paper, it is not important which mining industry determined a country's monetary policies so long as its money supply followed an exogenous process like equation (6) with $\gamma = 0$.

3.2 Unit-Root Tests

Before turning to our estimates of equations (12) and (13), we investigate whether our measures of p_t are difference-stationary and whether our measures of Δp_t and i_t are covariance stationary as our theory indicates. To that end, Table 1 reports the Dickey-Fuller test statistics for the series under consideration. The sample periods employed in calculating the statistics are the longest feasible ones.

The Dickey-Fuller statistics reject that our measures of Δp_t and i_t are difference stationary and except for Germany fail to reject that p_t is difference stationary. Failure to reject the null hypothesis of difference-stationarity for p_t does not mean that p_t is difference-stationary, however. We therefore examined our five series for p_t using the KPSS test, which makes covariance-stationarity the null hypothesis. Table 2 reports the test statistics that we obtained. We reject the null hypothesis of stationarity for all five series. On the whole, then, the evidence indicates that p_t is difference-stationary and i_t is covariance-stationary for each of the countries.

3.3 Results from Estimation

We used OLS to fit equations (12) and (13) to data for Belgium, France, Germany, the Netherlands, Switzerland, the United Kingdom and the United States, which were described in section 3.1. The resulting estimates are reported in Table 3. The Fama-Fisher effect is always estimated to be negative and often statistically significantly so¹⁵. By contrast, the Gibson effect is always estimated to be positive and highly significantly so.

These two findings are broadly consistent with our theory. Furthermore, the Gibson effect, which the theory implies should equal the interest semielasticity of money demand, is estimated to range between 2 and 4.5 years, a plausible value for this

¹⁵ Barsky and Bradford de Long (1991, p. 827) also found a negative Fama-Fisher effect for the United States and the United Kingdom during this period. They call this finding “the principal strike against the Fisher effect.”

Table 1
Augmented Dickey-Fuller Test Statistics

Series	Statistics
Logarithm of the Belgian Wholesale Price Index	-2.19 (0)
Logarithm of the French Wholesale Price Index	-2.82 (1)
Logarithm of the German Wholesale Price Index	-3.60 ^b (1)
Logarithm of the UK Wholesale Price Index	-1.84 (0)
Logarithm of the US Wholesale Price Index	-0.96 (0)
Differenced Logarithm of the Belgian Wholesale Price Index	-7.24 ^a (0)
Differenced Logarithm of the French Wholesale Price Index	-5.16 ^a (0)
Differenced Logarithm of the German Wholesale Price Index	-5.56 ^a (1)
Differenced Logarithm of the UK Wholesale Price Index	-8.28 ^a (0)
Differenced Logarithm of the US Wholesale Price Index	-5.44 ^a (0)
Belgian Free-Market Rate of Discount	-3.44 ^b (0)
French Open-Market Rate of Discount	-3.23 ^b (0)
German Open-Market Rate of Discount	-3.14 ^b (0)
Market Rate of Discount in Amsterdam	-4.10 ^a (0)
Average Swiss Discount Rate at Various Banks of Issue	-3.81 ^a (0)
UK Open-Market Rate of Discount	-5.32 ^a (0)
US Commercial Paper Rate	-5.97 ^a (0)

Notes. The integers in parentheses are the degrees of augmentation, which were chosen to minimize the Schwartz Information Criterion (SIC). The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively. The regressions for the interest rates and inflation rates include intercepts while those for the wholesale price indices include both intercepts and linear time trends, thereby allowing for the possibility that the price levels could trend under the alternative hypothesis of covariance-stationarity.

Table 2
KPSS Test Statistics

Series	Statistics
Logarithm of the Belgian Wholesale Price Index	2.21 ^a (0)
Logarithm of the French Wholesale Price Index	1.61 ^a (0)
Logarithm of the German Wholesale Price Index	0.72 ^a (1)
Logarithm of the UK Wholesale Price Index	7.52 ^a (0)
Logarithm of the US Wholesale Price Index	4.30 ^a (0)

Notes. The prewhitening filter for the test was based on spectral OLS with a lag length determined by the Schwartz Information Criterion (SIC). The lag length is indicated by the integer in parentheses. The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

Table 3
Annual Estimates of the Fama-Fisher and Gibson Effects
for Seven Countries over Various Pre-WWI Periods

Country and Sample Period	Estimate of β_F	Estimate of β_G
Belgium, 1849-1913	-0.29 (0.59)	2.35 ^a (0.76)
France, 1873-1913	-1.66 ^c (0.90)	4.40 ^a (1.10)
Germany, 1872-1913	-1.26 (1.09)	3.48 ^a (0.99)
Netherlands, 1874-1912	-1.41 (0.85)	2.59 ^a (0.76)
Switzerland, 1851-1913	-1.19 (0.82)	2.92 ^a (0.87)
United Kingdom, 1825-1913	-1.20 ^b (0.51)	2.72 ^a (0.43)
United States, 1879-1913	-2.86 ^a (1.04)	2.05 ^a (0.71)

Notes. The estimated regressions take the form $\Delta p_t = \alpha_F + \beta_F i_{t-1} + \varepsilon_{Ft}$ and $\Delta p_t = \alpha_G + \beta_G \Delta i_t + \varepsilon_{Gt}$, and the figures in parentheses are standard errors. The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

parameter.¹⁶

It is instructive to also estimate the contemporaneous Fisherian relationships

$$\Delta p_t = \phi_0 + \phi_1 i_t + \omega_t \quad (19)$$

and

$$i_t = \varphi_0 + \varphi_1 \Delta p_t + \bar{\omega}_t, \quad (20)$$

where the ϕ s and φ s are parameters and ω_t and $\bar{\omega}_t$ are error terms. According to our theory, the OLS estimators of ϕ_1 and φ_1 should converge in probability to $\lambda(1-\rho)$ and

¹⁶ For example, R. W. Hafer and Dennis Jansen (1991) estimated the long-run elasticity of demand for M2 with respect to the commercial paper rate in the United States over the period 1915-1988 to be 0.12, which implies an interest semielasticity of 3 years at a nominal interest rate of .04 per annum.

$\lambda^{-1} / \{2 + (1 - \rho)^{-1} \lambda^{-2} [1 + \lambda(1 - \rho)]^2 [\text{var}(\mu_i - \eta v_i) / \text{var} u_i]\}$, respectively.¹⁷ Therefore, when we estimate equation (19), we should obtain the negative of what we did when we fitted equation (12), and when we estimate equation (20), we should obtain a positive value that is strictly less than half the reciprocal of what we obtained when we fitted equation (13).

Table 4 reports the estimates obtained by fitting equations (19) and (20) to our data and repeats the estimates reported in Table 3 for easy comparison. Our estimates of ϕ_1 are indeed approximately the negatives of our estimates of the Fama-Fisher effect,¹⁸ and our estimates of φ_1 are indeed positive and strictly less than half the reciprocals of our estimates of the Gibson effect.

Traditional tests of Fisherian interest theory were based on OLS estimates of equation (20) or more elaborate version with lags of the inflation rate. Our theory provides an explanation for why such an approach is doomed to estimate values much less than one if prices are flexible and money is supplied inelastically.¹⁹ The probability limit of such estimators is strictly less than half the reciprocal of the interest semielasticity of money demand. If the semielasticity is 3 years as suggested in footnote 15 and the variance of aggregate-supply shocks, money-supply shocks and money-demand shocks is large relative to the variance of aggregate demand shocks, this probability limit may well lie between 0.025 and 0.064 years⁻¹, the range of estimates reported in Table 4.

The data that we have used to estimate the Fama-Fisher effect are inappropriate to some extent. Ideally, p_t and p_{t-1} in the regression

$$\Delta p_t = \alpha_F + \beta_F i_{t-1} + \varepsilon_{Ft}$$

should be the logarithms of the price level at the discrete times t and $t-1$ and i_{t-1} should be the yield at the discrete time $t-1$ on a bond maturing at the discrete time t . By contrast, our data are annual averages, and the terms for the interest rates are appreciably less than one

¹⁷ See the appendix for derivations.

¹⁸ In Table 4 the average of the estimates of ϕ_1 is 1.36, while the average of the estimates of β_F is -1.41.

¹⁹ Footnote 9 explains why “testing Fisherian interest theory” is impossible in any case.

Table 4
Comparison of Estimates from Contemporaneous Fisherian
Regressions with the Estimates in Table 3

Country Sample Period	ϕ_1 Estimate	β_F Estimate	ϕ_1 Estimate	β_G Estimate
Belgium 1849-1913	0.99 (0.59)	-0.29 (0.59)	0.043 (0.026)	2.35 ^a (0.76)
France 1873-1913	0.98 (0.89)	-1.65 ^c (0.90)	0.030 (0.027)	4.40 ^a (1.10)
Germany 1872-1913	2.06 ^b (1.02)	-1.26 (1.09)	0.044 ^b (0.022)	3.48 ^a (0.99)
Netherlands 1874-1912	1.43 ^c (0.82)	-1.41 (0.85)	0.052 ^c (0.030)	2.59 ^a (0.76)
Switzerland 1851-1913	1.08 (0.82)	-1.19 (0.82)	0.025 (0.019)	2.92 ^a (0.87)
United Kingdom 1825-1913	1.50 ^a (0.49)	-1.20 ^a (0.51)	0.064 ^a (0.021)	2.71 ^a (0.43)
United States 1879-1913	1.46 (1.11)	-2.84 ^a (1.04)	0.034 (0.026)	2.03 ^a (0.71)

Notes. For the estimates of β_F and β_G , see the notes in Table 3. The estimates of ϕ_1 and ϕ_1 are obtained by applying OLS to regressions of the form $\Delta p_t = \phi_0 + \phi_1 i_t + \omega_t$ and $i_t = \phi_0 + \phi_1 \Delta p_t + \omega_t$, and the figures in parentheses are standard errors. The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

year. For most of the countries, it is not possible to improve on the data that we have employed. For the United Kingdom and the United States, however, we can more closely approximate the theoretically appropriate series since we have monthly data. Specifically, let t now be a continuous index, and let i_t be the continuously compounded riskless nominal interest rate at time t with a term of n years to maturity. Then to a close approximation, the Fama-Fisher effect can be estimated by applying OLS to a regression of the form²⁰

$$n^{-1} \Delta_n p_t = \beta_F i_{t-n} + \varepsilon_{Ft}, \quad (12')$$

where $\Delta_n p_t \equiv p_t - p_{t-n}$

²⁰ Note that multiplying by the factor n^{-1} is necessary to annualize $\Delta_n p_t$.

Since UK open-market paper had a term of three months during the sample period, we estimate equation (12') with $n = 1/4$ and use data for March, June, September and December for a sample period from June 1824 to December 1913,²¹ obtaining the estimated regression

$$4\Delta_{1/4}\hat{p}_t = \frac{0.048}{(0.017)} - \frac{1.551}{(0.486)} i_{t-1/4}$$

where the numbers in parentheses are standard errors. The negative Fama-Fisher effect has strengthened and become more significant with improved data.

During the sample periods that we consider, US commercial paper had terms to maturity ranging from four to six months. On the assumption that each of these terms is equally represented in the data, equation (12') implies that

$$\nabla p_t = \alpha_F + \beta_F i_{t-1/2} + \bar{\varepsilon}_{Ft}, \quad (21)$$

where $\nabla p_t = \frac{1}{3}[3(p_{t-1/6} - p_{t-1/2}) + (12/5)(p_{t-1/12} - p_{t-1/2}) + 2(p_t - p_{t-1/2})]$ and

$\bar{\varepsilon}_{Ft} \equiv \frac{1}{3}[\varepsilon_{Ft-1/3} + \varepsilon_{Ft-5/12} + \varepsilon_{Ft-1/2}]$ Our assumptions imply that $\bar{\varepsilon}_{Ft}$ is orthogonal to $i_{t-1/2}$

and serially uncorrelated if the observations are spaced at least six months apart.

We used OLS to fit equation (21) to semiannual observations in June and December over the sample period extending from June 1879 to December 1913, obtaining

$$\nabla\hat{p}_t = \frac{0.152}{(0.047)} - \frac{2.965}{(0.953)} i_{t-1/12},$$

²¹ The monthly open-market discount rate for the United Kingdom is included in the NBER Macro History Database, which is available on NBER's website. However, NBER's collection on the monthly wholesale price index starts only in January 1885. From January 1824 to December 1850, we use an index from Arthur Gayer, Walt Rostow, and Anna Schwartz (1953, v. 1, pp. 468-470). Finally, from January 1851 to December 1885 we use an index from Jan Tore Klovland (1993), who interpolated the annual index of A. Sauerbeck (1886) into a monthly index. No adjustment is needed between the series of Klovland (1993) and NBER. A five-year overlap between the series of Gayer, Rostow, and Schwartz and Klovland (1993) is used to adjust the level of the former series. It is possible to triple the sample size by employing every monthly observation on $\Delta_{1/4}p_t$ and $i_{t-1/4}$, but doing so would require accounting for the induced monthly noninvertible second-order moving-average structure of the error term. Spacing the observations so that the terms of the open-market paper do not overlap avoids this problem, albeit with some loss of efficiency.

where again the numbers in parentheses are standard errors.²² Once again, better data make the estimated Fama-Fisher effect even more negative, and it remains statistically significant at the 0.01 level.

4 Empirical Analysis for the Interwar and Postwar Periods

Our model provides new insights for periods after World War I. Central banks took many years to learn how to be effective. The learning process may not immediately have given rise to “elastic currencies;” i.e., large values of γ . Furthermore, prices may have remained fairly flexible for an appreciable time. As a result, the Fama-Fisher effect may have remained negative for many years, and a positive Gibson effect may also have been present. Even when monetary policy began to take the form of interest-rate control and became focused on domestic objectives and prices became distinctly sticky, the Gibson effect may still have remained positive because theory does not rule out this possibility. For these reasons, we estimate Fama-Fisher and Gibson effects using data for the interwar and postwar periods.

4.1 The Interwar Period

Table 5 reports interwar estimates of the Fama-Fisher and Gibson effects for France, the United Kingdom, and the United States. The French and UK data are quarterly for March, June, September and December,²³ while the US data are semiannual observations for June and December.²⁴ The parameter β_F is estimated from regressions of the form (12') and (21).

²² The data come from Macaulay (1938) and Warren and Pearson (1935).

²³ Both the French and UK interest rates are open-market discount rates. The French price level is measured by the Paris Retail Price Index for 34 articles, while the UK price level continues to be measured by the wholesale price index for all commodities. All series come from the NBER Macro History Database.

²⁴ Our interest rate for the United States is the commercial paper rate, which we obtained from *Banking and Monetary Statistics, 1914-1941*. We measured the US price level with the consumer price index, which we obtained in a private communication from the Bureau of Labor Statistics. It is also available from the NBER Macro History Database. When available we prefer to use the consumer price index in preference to the wholesale price index since it better approximates the concept of the price level in our theory.

Table 5

**Estimates of the Fama-Fisher and Gibson Effects
for France, the United States, and the United Kingdom
in the Interwar Period**

Sample Period	Estimate of β_F	Estimate of β_G
France, 1929Q2-1939Q3	4.80 ^a (1.74)	-0.20 (0.50)
United Kingdom, 1919Q2-1939Q2	-3.25 ^a (1.15)	1.59 ^b (0.75)
United Kingdom, 1919Q2-1929Q4	-9.39 ^a (2.70)	2.94 ^b (1.42)
United Kingdom, 1930Q1-1939Q2	-3.09 (2.01)	0.90 (0.77)
United States, 1919H2-1941H2	-0.54 (0.74)	2.83 ^b (1.11)
United States, 1919H2-1929H2	-2.50 (1.62)	3.11 ^a (1.05)
United States, 1930H1-1941H2	-3.64 ^c (1.88)	2.44 (2.11)

Notes. The data for France and the United Kingdom are quarterly, and the estimated regressions take the form $4\Delta_{1/4}p_t = \alpha_F + \beta_F i_{t-1/4} + \varepsilon_{Ft}$ and $\Delta_{1/4}p_t = \alpha_G + \beta_G \Delta_{1/4}i_t + \varepsilon_{Gt}$. The data for the United States are semiannual, and the estimated regressions take the form $\nabla p_t = \alpha_F + \beta_F i_{t-1/2} + \bar{\varepsilon}_{Ft}$ and $\Delta_{1/2}p_t = \alpha_G + \beta_G \Delta_{1/2}i_t + \varepsilon_{Gt}$. The figures in parentheses are standard errors, and the superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

For the United Kingdom and the United States, the Fama-Fisher effect is estimated to be negative and the Gibson effect is estimated to be positive in all sample periods, and both effects are often statistically significant. France is very different: its Fama-Fisher effect is estimated to be significantly *positive* and its Gibson effect is estimated to be negative, albeit not significantly so.

Table 6 reports interwar estimates of the Fama-Fisher and Gibson effects for three other European countries for which only annual data are available. It is based on annual average data from Tables 65, 67 and 71 of Homer and Sylla (1991), and Table II of Mitchell (1992). The Fama-Fisher effect is estimated to be positive and statistically insignificant for Belgium but negative and statistically significant for the Netherlands and

Table 6

Annual Estimates of the Fama-Fisher and Gibson Effects for Selected European Countries in the Interwar Period

Country and Sample Period	Estimate of β_F	Estimate of β_G
Belgium, 1922-1940	2.58 (2.35)	2.16 (4.05)
Netherlands, 1920-1940	-3.02 ^c (1.69)	1.85 (2.29)
Switzerland, 1922-1940	-4.28 ^c (2.42)	4.49 (2.84)

Notes. The estimated regressions take the form $\Delta p_t = \alpha_F + \beta_F i_{t-1} + \varepsilon_{Ft}$ and $\Delta p_t = \alpha_G + \beta_G \Delta i_t + \varepsilon_{Gt}$, and the figures in parentheses are standard errors. The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

Switzerland. The Gibson effect is estimated to be positive for all of the countries but is not statistically significant for any.

On the whole, Tables 5 and 6 indicate that the Gibson's Paradox was alive and well in the interwar period and indeed similar to the one observed before World War I. That it continued to hold indicates that *it is not merely a gold- or commodity-standard phenomenon*. Rather, it arises under the more general conditions described in our model. Moreover, a negative Fama-Fisher effect is always accompanied by a positive Gibson effect, a feature clearly captured by case I of our model.

Keynes and many other observers believed that the Gibson Paradox no longer held after World War I. They can be excused for this belief, which in most cases was based on eyeball econometrics applied to levels of the series. Because of major discoveries in time-series econometrics in recent years, we now know that the relationship between a differenced-stationary and a covariance-stationary series cannot be so readily assessed.

Table 7

**Quarterly Estimates of the Fama-Fisher and Gibson Effects
for the United States in the Postwar Period**

Sample Period	Estimate of β_F	Estimate of β_G
1953Q2-2004Q2	0.71 ^a (0.07)	0.08 (0.05)
1953Q2-1960Q4	0.84 ^b (0.34)	-0.12 (0.15)
1961Q1-1979Q4	1.60 ^a (0.12)	0.29 ^c (0.15)
1980Q1-2004Q2	0.53 ^a (0.08)	0.04 (0.06)

Notes. The estimated regressions take the form $4\Delta_{1/4}p_t = \alpha_F + \beta_F i_{t-1/4} + \varepsilon_{Ft}$ and $\Delta_{1/4}p_t = \alpha_G + \beta_G \Delta_{1/4}i_t + \varepsilon_{Gt}$, and the figures in parentheses are standard errors. The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively. The sample period begins in 1953Q2 in order to avoid the Pre-Accord Period.

4.2 The Postwar Period

Table 7 reports postwar estimates of the Fama-Fisher and Gibson effects for the United States. It is based on quarterly observations in March, June, September and December on the three-month Treasury bill rate and the consumer price index.²⁵ The Fama-Fisher effect is estimated to be positive and statistically significant, while the Gibson effect is typically estimated to be small and statistically insignificant. In keeping with the findings of Romer and Romer (2002) and Clarida, Galí and Gertler (2000), the Fama-Fisher effect is estimated to be less than one over the sample periods 1953Q2-1960Q4 and 1980Q1-2004Q2 and to be greater than one in the sample period 1961Q1-1979Q4.^{26,27}

²⁵ We downloaded the data on the three-month Treasury bill rate from the website of the US Federal Reserve Board and the data on the consumer price index from the website of the US Bureau of Labor Statistics. The latter series was seasonally adjusted. These data and all others used in this study are available from either of the authors upon request.

²⁶ Clarida, Galí and Gertler (2000) estimated that τ appreciably exceeded one during the Volcker/Greenspan monetary regime (1980Q1-present) and was somewhat less than one during the previous monetary regime (1960Q1-1979Q4). Romer and Romer (2002) estimated that it exceeded one in the monetary regime in place during the 1950s.

²⁷ Fama fitted a similar equation over a sample period from January 1953 to July 1971, obtaining an estimate of β_F that did not differ significantly from one. He interpreted this finding as providing evidence

Table 8

Quarterly Estimates of the Fama-Fisher and Gibson Effects for Selected Developed Countries in the Postwar Period

Country and Sample Period	Estimate of β_F	Estimate of β_G
Belgium, 1957Q2-2003Q4	0.69 ^a (0.26)	0.24 (0.29)
Canada, 1957Q2-2003Q4	0.54 ^a (0.06)	0.11 ^c (0.06)
France, 1970Q2-2002Q4	0.78 ^a (0.08)	0.16 ^b (0.08)
Germany, 1953Q2-1990Q2	0.56 ^a (0.12)	0.03 (0.07)
Italy, 1977Q2-2003Q4	0.78 ^a (0.07)	0.16 (0.12)
Japan, 1981Q1-2002Q2	0.36 ^a (0.09)	0.05 (0.13)
Netherlands, 1962Q2-1998Q4	0.38 ^b (0.18)	-0.03 (0.05)
Switzerland, 1980Q2-2003Q4	0.63 ^a (0.10)	0.07 (0.10)
United Kingdom, 1964Q2-2003Q4	0.92 ^a (0.13)	0.08 (0.11)

Notes. Except for Germany and the Netherlands, the estimated regressions take the form $4\Delta_{1/4}p_t = \alpha_F + \beta_F i_{t-1/4} + \varepsilon_{Ft}$ and $\Delta_{1/4}p_t = \alpha_G + \beta_G \Delta i_t + \varepsilon_{Gt}$. To obtain a consistent estimate of the Fama-Fisher effect for those countries where we are constrained to use the call money rate, we employ the slightly different regression equation $4\Delta_{1/4}p_t = \alpha_F + \beta_F \bar{c}_t + \tilde{\varepsilon}_{Ft}$, where \bar{c}_t is the average call money rate over quarter t . This regression equation must be estimated using instrumental variables dated no later than two quarters before date t ; we use $\bar{c}_{t-1/2}$ for this purpose. (See the appendix for the details.) The figures in parentheses are standard errors. The superscripts a, b, and c indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

Table 8 reports postwar estimates of the Fama-Fisher and Gibson effects for nine other countries. For Belgium, Canada, France, Italy, Switzerland and the United Kingdom, it is based on quarterly observations in March, June, September, and December of the

for his joint hypothesis of rational expectations and a constant expected real interest rate. When we estimate over the sample period 1953Q2-1971Q2, we obtain an estimate that also differs insignificantly from one. We interpret this finding very differently, however. Fama's sample period spans a period in which τ exceeded one (1953Q2-1959Q4) and another period when it fell short of one (1960Q1-1971Q2). By happenstance, Fama's estimate just turned out to be too close to one to reject the null hypothesis of $\beta_F = 1$.

country's three-month Treasury bill rate and its consumer price index.²⁸ We must use a three-month interest rate on private bills for Japan, while for Germany and the Netherlands, we must use the call money rate since a Treasury bill rate is not available continuously.²⁹

The Fama-Fisher effect is estimated to be highly significantly positive for every country and also less than one as our theory indicates. The Gibson effect is estimated to be positive for all but one country though appreciably smaller than we found for the periods before World War II. Although the Gibson effect is still mostly estimated to be positive, Gibson's Paradox is not a pervasive feature of postwar economies, while a strong positive Fama-Fisher effect is. The most likely reason for this juxtaposition of effects is that monetary policy now provides an infinitely "elastic currency" in the form of Taylor-type interest-rate policies. The postwar period probably also has appreciably stickier prices than earlier periods.

5 Conclusions

This paper specifies a set of conditions under which the Fama-Fisher effect is negative and the Gibson effect is positive, a conjunction that is doubly paradoxical. Unlike Barsky and Summers (1988), our model is not anchored to the classical gold standard. Instead, we find that the presence of Gibson's Paradox depends crucially on price flexibility and the elasticity of the currency. In particular, when prices are flexible and money is inelastically supplied, a negative Fama-Fisher effect and a positive Gibson effect are simply two sides of the same coin.

Of course, the classical gold-standard period does indeed satisfy both conditions and as a result Gibson's Paradox does characterize that period. Surprisingly, however, Gibson's

²⁸ All of the data come from the CD-Rom for October 2004 of the *International Financial Statistics* of the International Monetary Fund.

²⁹ German and Dutch call money rates come from monthly hardcopy issues of the *International Financial Statistics* of the International Monetary Fund. Using call money rates creates some econometric problems. The appendix develops adjustments that can approximately handle these problems.

Paradox also largely characterized the interwar period, a period in which the classical gold standard no longer prevailed. Our theory has a ready explanation for this phenomenon: currencies had not yet become elastic and prices continued to be fairly flexible. By contrast, Gibson’s Paradox is not a pervasive feature of the postwar period. A strong positive Fama-Fisher effect is evident and the Gibson effect is only weakly positive—if positive at all. In our theory, this combination of effects can arise if the central bank pursues a Taylor-type interest-rate rule and no doubt more generally when money is elastically supplied. After World War II, central banks in many developed countries became more active in accommodating liquidity demand and more focused on domestic objectives than they had been during earlier periods.³⁰ In the meantime, many of them also adopted aggressive inflation-targeting rules. All of these institutional changes have had a profound impact on the dynamics of the actual and expected inflation rate. With sticky prices, inflation can be readily forecasted, enabling an inflation-targeting central bank to induce a positive Fama-Fisher effect. At the same time, the price level and interest rate no longer bears any necessary relationship with each other, thus making the Gibson effect indeterminate in sign.

6 Appendix

6.1 Introducing Persistency into e_t

In this section we modify the assumption that e_t is an i.i.d. white noise. Instead, we assume that it follows the first-order autoregression process $e_t = \theta e_{t-1} + \xi_t$, where $|\theta| < 1$ and ξ_t is an i.i.d. white noise with finite variance.

In this case, the minimum-state-variable solution to equation (15) is

$$\begin{aligned} x_t = & \left[(1-\rho)(1-\beta\rho) + \rho\varphi(\tau-1)\kappa \right]^{-1} (1-\beta\rho)\varphi u_t \\ & - \left[(1-\theta)(1-\beta\theta) + \theta\varphi(\tau-1)\kappa \right]^{-1} (1-\beta\theta)\varphi e_t \end{aligned} \tag{16'}$$

³⁰ A commonly known critique of central banks during the Great Depression is their failure to supply enough liquid assets; e.g., Milton Friedman and Anna Jacobson Schwartz (1963) and Charles Kindleberger (1973).

Substituting (16') into (14), using (3) and $e_t = \theta e_{t-1} + \xi_t$, and rearranging, we obtain

$$\Delta p_t = Au_t - Be_t \quad (17')$$

where

$$A \equiv [(1-\rho)(1-\beta\rho)\varphi^{-1}\kappa^{-1} + \rho(\tau-1)]^{-1}$$

$$B \equiv [(1-\theta)(1-\beta\theta)\varphi^{-1}\kappa^{-1} + \theta(\tau-1)]^{-1}.$$

Substituting (17') into (7), we have

$$i_t = A\tau\rho u_t + (1-B\theta\tau)e_t. \quad (18')$$

The OLS estimator of β_F converges in probability to

$$\begin{aligned} \frac{\text{cov}(\Delta p_{t+1}, i_t)}{\text{var}(i_t)} &= \frac{A^2\rho\tau \text{cov}(u_{t+1}, u_t) - B(1-B\theta\tau) \text{cov}(e_{t+1}, e_t)}{A^2\rho^2\tau^2 \text{var} u_t + (1-B\theta\tau)^2 \text{var} e_t} \\ &= \frac{A^2\rho^2\tau \text{var} u_t - B\theta(1-B\theta\tau) \text{var} e_t}{A^2\rho^2\tau^2 \text{var} u_t + (1-B\theta\tau)^2 \text{var} e_t} \\ &= \tau^{-1} \left[\frac{(1-\theta^2)A^2\rho^2\tau^2 \text{var} u_t - B\theta\tau(1-B\theta\tau) \text{var} \xi_t}{(1-\theta^2)A^2\rho^2\tau^2 \text{var} u_t + (1-B\theta\tau)^2 \text{var} \xi_t} \right] \end{aligned}$$

since $\text{var} e_t = \text{var} \xi_t / (1-\theta^2)$.

If $1-B\theta\tau > 0$, the value in braces is strictly less than 1, and can even be negative if $\text{var} \xi_t \gg \text{var} u_t$. If $1-B\theta\tau < 0$, the value in braces is strictly greater than 1. Finally, if $1-B\theta\tau = 0$, the value in braces equals 1.

It is easy to verify that when $\theta = 0$, $\text{var} e_t = \text{var} \xi_t$ and the probability limit is the same as that derived in Case II. When θ approaches 1, B approaches $(\tau-1)^{-1}$ and the probability limit approaches 1.

Consider the benchmark case of $1-B\theta\tau = 0$. Substituting $1/\theta\tau$ into the definition of B and rearranging, we obtain the following quadratic equation for θ .

$$f(\theta) \equiv \beta\theta^2 - (1 + \beta + \varphi\kappa)\theta + 1 = 0$$

The function $f(\theta)$ is a parabola that opens upward and has an intercept on the ordinate of 1. Since $f(1) = -\varphi\kappa < 0$, it also has two positive roots, one on each side of 1. Denote the smaller root as θ_1 . Since $f'(1) = 2\beta - (1 + \beta + \varphi\kappa) < 0$, the minimum of $f(\theta)$ is to the right of $\theta = 1$. This function is shown in Figure A. Note that when θ lies on $(0, \theta_1)$, there is no lower bound for the probability limit and it might even be negative. On the other hand, when θ lies on $(\theta_1, 1)$, there is no upper bound for the probability limit. These results are summarized in Table A. When θ is close to θ_1 , however, the probability limit is τ^{-1} . For this reason, we regard β_F as informative about the parameter τ and hence about how monetary policy was conducted; hence, our analysis for the postwar period.

The OLS estimator of β_G converges in probability to

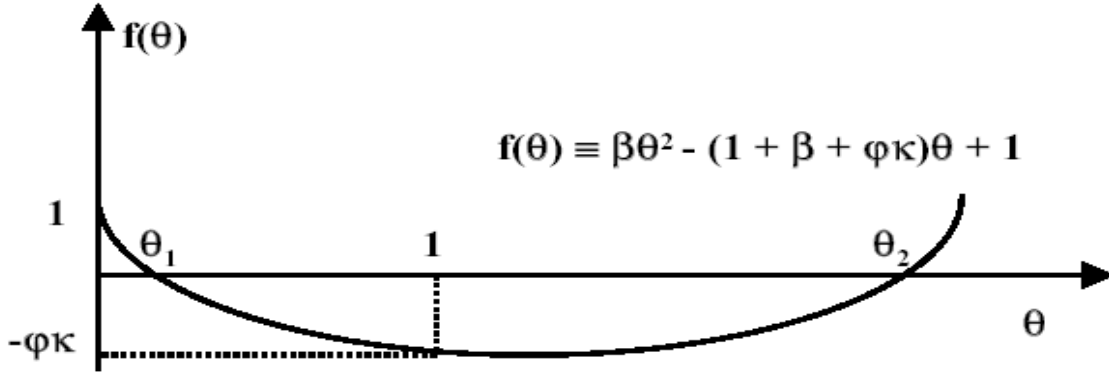
$$\begin{aligned} \frac{\text{cov}(\Delta p_t, \Delta i_t)}{\text{var}(\Delta i_t)} &= \frac{A^2 \rho \tau \text{cov}(u_t, \Delta u_t) - B(1 - B\theta\tau) \text{cov}(e_t, \Delta e_t)}{A^2 \rho^2 \tau^2 \text{var}(\Delta u_t) + (1 - B\theta\tau)^2 \text{var}(\Delta e_t)} \\ &= \frac{1}{2} \left[\frac{A^2 \rho \tau (1 - \rho) \text{var } u_t - B(1 - B\theta\tau)(1 - \theta) \text{var } e_t}{A^2 \rho^2 \tau^2 (1 - \rho) \text{var } u_t + (1 - B\theta\tau)^2 (1 - \theta) \text{var } e_t} \right] \\ &= \frac{1}{2} \rho^{-1} \tau^{-1} \left[\frac{(1 + \theta) A^2 \rho^2 \tau^2 (1 - \rho) \text{var } u_t - B \rho \tau (1 - B\theta\tau) \text{var } \varepsilon_t}{(1 + \theta) A^2 \rho^2 \tau^2 (1 - \rho) \text{var } u_t + (1 - B\theta\tau)^2 \text{var } \varepsilon_t} \right] \end{aligned}$$

It is easy to verify that when $\theta = 0$, this is the same as that derived in Case II. The more general result is summarized in Table A below.

Table A. Dependence of Probability Limits on θ

θ	$\text{plim } \beta_F$	$\text{plim } \beta_G$
0	$(0, \tau^{-1})$	$(-\infty, \frac{1}{2}\rho^{-1} \tau^{-1})$
$(0, \theta_1)$	$(-\infty, \tau^{-1})$	$(-\infty, \frac{1}{2}\rho^{-1} \tau^{-1})$
θ_1	τ^{-1}	$\frac{1}{2}\rho^{-1} \tau^{-1}$
$(\theta_1, 1)$	$(\tau^{-1}, +\infty)$	$(0, +\infty)$
1	1	$(0, +\infty)$

Figure A



6.2 Probability Limits for the Contemporaneous Regressions

In case I, the OLS estimator of ϕ_1 converges in probability to

$$\frac{\text{cov}(\Delta p_t, i_t)}{\text{var } i_t} = \frac{[1 + \lambda(1 - \rho)]^{-2} \lambda \text{cov}(\Delta u_t, u_t)}{[1 + \lambda(1 - \rho)]^{-2} \text{var } u_t} = \lambda(1 - \rho) > 0.$$

The OLS estimator of ϕ_1 converges in probability to

$$\begin{aligned} \frac{\text{cov}(i_t, \Delta p_t)}{\text{var}(\Delta p_t)} &= \frac{[1 + \lambda(1 - \rho)]^{-2} \lambda \text{cov}(u_t, \Delta u_t)}{[1 + \lambda(1 - \rho)]^{-2} \lambda^2 \text{var}(\Delta u_t) + \text{var}(\mu_t - \eta v_t)} \\ &= \frac{[1 + \lambda(1 - \rho)]^{-2} \lambda(1 - \rho) \text{var } u_t}{[1 + \lambda(1 - \rho)]^{-2} 2\lambda^2(1 - \rho) \text{var } u_t + \text{var}(\mu_t - \eta v_t)} \\ &= \frac{\lambda^{-1}}{2 + [1 + \lambda(1 - \rho)]^2 \lambda^{-2}(1 - \rho)^{-1} [\text{var}(\mu_t - \eta v_t) / \text{var } u_t]} \end{aligned}$$

which is strictly positive and smaller than $\frac{1}{2}\lambda^{-1}$. Moreover, it is smaller, the larger $\text{var}(\mu_t - \eta v_t) / \text{var } u_t$ is.

In case II, the OLS estimator of ϕ_1 converges in probability to

$$\frac{\text{cov}(\Delta p_t, i_t)}{\text{var } i_t} = \frac{[(1 - \rho)(1 - \beta\rho)\kappa^{-1} + \rho\varphi(\tau - 1)]^{-2} \varphi^2 \tau \rho \text{var } u_t - \kappa\varphi \text{var } e_t}{[(1 - \rho)(1 - \beta\rho)\kappa^{-1} + \rho\varphi(\tau - 1)]^{-2} \varphi^2 \tau^2 \rho^2 \text{var } u_t + \text{var } e_t}$$

which is strictly less than $\tau^{-1}\rho^{-1}$ and can even be negative if $\text{var } e_t \gg \text{var } u_t$.

Finally, the OLS estimator of φ_1 converges in probability to

$$\frac{\text{cov}(i_t, \Delta p_t)}{\text{var}(\Delta p_t)} = \frac{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \tau \rho \text{var } u_t - \kappa\varphi \text{var } e_t}{\left[(1-\rho)(1-\beta\rho)\kappa^{-1} + \rho\varphi(\tau-1) \right]^{-2} \varphi^2 \text{var } u_t + \kappa^2 \varphi^2 \text{var } e_t}$$

which is strictly less than $\tau\rho$ and can even be negative if $\text{var } e_t \gg \text{var } u_t$.

6.3 Adjustment for Using the Call Money Rate

Let t be continuous and c_t be the call money rate at time t . For simplicity, call money is assumed to have a maturity of zero. From the definition of the assorted premia below, we have

$$\begin{aligned} i_{t-1/4} &= E_{t-1/4} \left(\int_0^{1/4} c_{t+s-1/4} ds / \int_0^{1/4} ds \right) + \left(\begin{array}{l} \text{term, risk and} \\ \text{liquidity premia} \end{array} \right) \\ &= E_{t-1/4} \bar{c}_t + \left(\begin{array}{l} \text{term, risk and} \\ \text{liquidity premia} \end{array} \right) \end{aligned} \quad (22)$$

where $\bar{c}_t \equiv 4 \int_0^{1/4} c_{t-s} ds$. It follows that

$$4\Delta_{1/4} p_t = \beta_F \bar{c}_t + \tilde{\varepsilon}_{Ft}, \quad (23)$$

where

$$\tilde{\varepsilon}_{Ft} \equiv \varepsilon_{Ft} - \beta_F (\bar{c}_t - E_{t-1/4} \bar{c}_t) + \beta_F \left(\begin{array}{l} \text{stochastic component} \\ \text{of the various premia} \end{array} \right). \quad (24)$$

According to the expectation theory of the term structure, the last term of equation (24) should be zero. More generally, it may have a variance that is small relative to the variance of the other two terms, or it may be approximately uncorrelated with information available on or before time $t-1/2$. In either case, $\tilde{\varepsilon}_{Ft}$ is approximately uncorrelated with information available on or before time $t-1/2$. Then to a close approximation, equation (23) can be estimated consistently by using the quarterly average of the call money rate lagged two quarters as an instrumental variable.

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