Fisher, Barro, and the Italian Interest Rate, 1845–93

Vito Antonio Muscatelli, University of Glasgow
Franco Spinelli, Universita’ di Brescia

1. INTRODUCTION

There is a growing international literature on the determinants of long-run interest rates in the OECD economies (see inter alia Friedman and Schwartz, 1982; Summers, 1983, 1986; Barro, 1987; Barsky, 1987; Evans, 1987; Mankiw, 1987; Boudoukh and Richardson, 1993). However, this literature has neglected the experience of Italy.1 Italy’s experience is of interest because it has been more inflation prone than the other G7 countries, and has suffered from recurring fiscal deficits. It is, therefore, a natural testing ground for some of the main theories of interest rate determination.

We focus on two issues in particular. First, we examine whether nominal interest rates have fully incorporated inflationary expectations (the “Fisher effect”) in an environment where inflation has been high and persistent. Second, we estimate the impact of fiscal policy on real interest rates in an attempt to verify the

Address correspondence to V. A. Muscatelli, University of Glasgow, Department of Political Economy, Adam Smith Bldg., Glasgow G12 8RT, UK.

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1 With the exception of Muscatelli and Spinelli (1996), who compare the evidence on the Gibson Paradox in the United States, United Kingdom, and Italy.
neoclassical model of fiscal policy. The latter issue is of particular interest, given the dramatic budgetary imbalances experienced by Italy over the last few decades. The period post-1970 warrants investigation because it represents a period of sustained fiscal expansion accompanied first by an accommodating monetary policy and, from the mid-1980s onwards, increasing monetary restraint.

Our key results are as follows. First, a significant Fisher effect is only detectable in the post-World War II period. Second, we find that temporary increases in government spending have a significant impact on real interest rates, as predicted by neoclassical theories of fiscal policy. However, debt is also found to affect the real interest rate in the late 1980s, which suggests that Ricardian equivalence does not hold.

This paper is structured as follows. In Section 2, we take a preliminary look at the behavior of interest rates in Italy over the last 150 years, and provide a brief outline of the main changes in fiscal and monetary policy over this period. In Sections 3 and 4 we present some formal econometric evidence on the behavior of interest rates in Italy. Tests of the Fisher effect are reported in Section 3, and the responsiveness of real interest rates to fiscal policy is examined in Section 4. Section 5 concludes.

2. INTEREST RATES AND PRICES IN THE LONG RUN

2A. A Graphical and Statistical Summary

Our interest rate data consists of a long-term rate, $R$ which is the yield (net of taxes) on the long-term government bonds for the whole sample 1845–1993. The price series used for our study is a cost-of-living index, which is consistent for the whole sample period. The series employed are reported in Spinelli and Fratianni (1991), and have been extended using national data sources. The difference between the long rate ($R$) and the rate of change of the price variable (II) is shown in Figure 1 for various subperiods, including the war years.

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1Our data refers to the unified Italian state post-1860, and to its immediate precursor, the Kingdom of Sardinia (Piedmont) for the period 1845–59.
2To avoid confusion later in the paper, we do not refer to $R$-II as the real interest rate, as this requires estimating expected inflation.
In the 19th century the long-term nominal interest rate fluctuated considerably. The interest rate was about 5.5 percent in 1845 and peaked to 9.5 percent in 1866, at the time of the Third War of Independence with Austria. Subsequently, the interest rate declined steadily to about 3.5 percent in 1913. The first World War marked the beginning of a very different story as the interest rate, through successive jumps, reached the values of 5.2 percent in 1930, 7.2 percent in 1957, and 20.2 percent in 1982. If we discount the temporary and partial declines of the 1930s and 1960s, on the whole, we have about 70 years of interest rate increases. Only in the post-1982 period does \( R \) appear to have inverted its upward trend, but it remains at levels much higher than those of the past.

As far as \( R-II \) is concerned (see Figure 1), there are two key points to note. First, nominal interest rates do not keep pace with the high inflations of World War I, World War II, and the 1970s. Second, from 1983 onwards, as inflation is slowly brought under control, the real interest rate remains fairly constant.
Table 1: Means and Standard Deviations of Interest Rates and Inflation for Various Subperiods of the Sample

<table>
<thead>
<tr>
<th>Date</th>
<th>R</th>
<th>II</th>
<th>R-II</th>
</tr>
</thead>
<tbody>
<tr>
<td>1845–1914</td>
<td>5.4</td>
<td>0.8</td>
<td>4.6</td>
</tr>
<tr>
<td></td>
<td>(1.5)</td>
<td>(4.5)</td>
<td>(4.6)</td>
</tr>
<tr>
<td>1845–60</td>
<td>5.9</td>
<td>1.1</td>
<td>4.8</td>
</tr>
<tr>
<td></td>
<td>(0.4)</td>
<td>(5.4)</td>
<td>(5.4)</td>
</tr>
<tr>
<td>1861–70</td>
<td>8.3</td>
<td>2.0</td>
<td>6.3</td>
</tr>
<tr>
<td></td>
<td>(1.0)</td>
<td>(5.9)</td>
<td>(6.1)</td>
</tr>
<tr>
<td>1871–80</td>
<td>5.9</td>
<td>1.9</td>
<td>4.0</td>
</tr>
<tr>
<td></td>
<td>(0.6)</td>
<td>(7.2)</td>
<td>(7.2)</td>
</tr>
<tr>
<td>1881–90</td>
<td>4.7</td>
<td>−0.6</td>
<td>5.2</td>
</tr>
<tr>
<td></td>
<td>(0.2)</td>
<td>(3.0)</td>
<td>(3.1)</td>
</tr>
<tr>
<td>1891–1900</td>
<td>4.4</td>
<td>−0.5</td>
<td>4.9</td>
</tr>
<tr>
<td></td>
<td>(0.3)</td>
<td>(0.8)</td>
<td>(1.0)</td>
</tr>
<tr>
<td>1901–10</td>
<td>3.9</td>
<td>1.0</td>
<td>2.9</td>
</tr>
<tr>
<td></td>
<td>(0.1)</td>
<td>(2.2)</td>
<td>(2.2)</td>
</tr>
<tr>
<td>1911–20</td>
<td>4.1</td>
<td>15.0</td>
<td>−10.9</td>
</tr>
<tr>
<td></td>
<td>(0.4)</td>
<td>(17.4)</td>
<td>(17.0)</td>
</tr>
<tr>
<td>1921–30</td>
<td>4.9</td>
<td>2.3</td>
<td>2.6</td>
</tr>
<tr>
<td></td>
<td>(0.3)</td>
<td>(8.5)</td>
<td>(8.6)</td>
</tr>
<tr>
<td>1931–40</td>
<td>5.0</td>
<td>2.4</td>
<td>2.6</td>
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<tr>
<td></td>
<td>(0.6)</td>
<td>(8.2)</td>
<td>(7.8)</td>
</tr>
<tr>
<td>1941–50</td>
<td>6.0</td>
<td>62.6</td>
<td>−56.6</td>
</tr>
<tr>
<td></td>
<td>(0.4)</td>
<td>(104.4)</td>
<td>(104.1)</td>
</tr>
<tr>
<td>1951–60</td>
<td>6.3</td>
<td>3.5</td>
<td>2.7</td>
</tr>
<tr>
<td></td>
<td>(0.6)</td>
<td>(2.7)</td>
<td>(2.6)</td>
</tr>
<tr>
<td>1961–70</td>
<td>5.7</td>
<td>3.9</td>
<td>1.8</td>
</tr>
<tr>
<td></td>
<td>(0.8)</td>
<td>(2.0)</td>
<td>(2.1)</td>
</tr>
<tr>
<td>1971–80</td>
<td>10.9</td>
<td>14.2</td>
<td>−3.2</td>
</tr>
<tr>
<td></td>
<td>(3.3)</td>
<td>(5.6)</td>
<td>(3.8)</td>
</tr>
<tr>
<td>1981–93</td>
<td>13.5</td>
<td>8.3</td>
<td>5.2</td>
</tr>
<tr>
<td></td>
<td>(3.6)</td>
<td>(4.2)</td>
<td>(0.9)</td>
</tr>
</tbody>
</table>

In Table 1 we report the means and standard deviations (in brackets) of the three key variables for various subsamples. The two World Wars and the 1970s are characterized by negative values of R-II. Up to 1913 and during the 1980s, the gap between nominal interest rates and inflation was positive, and often large by international standards.

To illustrate the forces that caused these major changes in interest rate behavior, we outline the main events that shaped Italian macroeconomic policy in the last 150 years.
2B. Monetary and Fiscal Regimes—A Brief Survey

Italian political unification took place gradually through a process of annexation and the acquisition of new territories on the part of the small Kingdom of Sardinia (consisting mainly of Piedmont and adjoining territories). The Kingdom of Sardinia became the Kingdom of Italy in 1860, and unification of the whole Italian peninsula was completed in 1970. Our sample partly covers a period prior to unification, 1845–59, where the data refers to the Kingdom of Sardinia. Post-1860 data refers to the Kingdom of Italy.

The creation of the first monetary issuing authorities in the Kingdom of Sardinia took place between 1844 and 1849. At first the issuing authorities, the Banca di Genova and the Banca di Torino4 (these were merged into a single issuing bank in 1849—the Banca Nazionale degli Stati Sardi) issued notes that were fully convertible in specie. However, the first difficulties in maintaining convertibility were encountered in 1848 when, following an increase in military expenditures, the issuing banks were authorized to enact a loan to the Treasury and to suspend convertibility until 1851.

The next suspension of convertibility came in 1859–60 during the next phase of the struggle for unification. During the 1860s the links between the issuing authorities and the Treasury were formalized, so that bank notes could be issued to meet the demands of the state over and above the issuing banks’ reserve requirements and legal maximum circulation limits.5 A further suspension of convertibility came in 1866. It was triggered by an expansionary fiscal policy due to the expenditures incurred in the war with Austria.

A slow process of fiscal retrenchment in the 1870s eventually allowed Italy to return to full convertibility in 1880. However,

4The account presented here is brief for reasons of space. A full history of the creation and merger of the various monetary issuing authorities in Italy can be found in Spinelli and Fratianni (1991).

5Another complication that hampered monetary control is that, following the gradual process of annexation of the small independent states into the Kingdom of Italy, some of the preunification states’ issuing authorities retained their independent status. The Banca Nazionale was only one of six competing issuing authorities in 1870. The number of issuing banks was reduced to only three in 1893 when the Banca Romana failed, and the Banca Nazionale took over its liabilities and those of two of the smaller issuing authorities and became the Banca d’Italia. Not until 1926 did the Banca d’Italia become the sole monetary issuing authority.
this return to the Gold Standard was short lived. Following a combination of domestic inflationary pressures and the drying up of capital inflows, the Lira came under pressure in the mid-1880s. What followed was a de facto suspension of convertibility from the mid-1880s, as the issuing authorities refused to convert paper money. This was eventually matched by a de jure suspension of convertibility in 1893. Monetary stability was slowly restored in the period 1894–1913, but without a formal return to the Gold Standard. De facto, exchange rates were stabilized during the period 1902–13, through more prudent fiscal and monetary policies.

Overall, the period 1845–1913 saw a very mixed participation to the Gold Standard, but this did not prevent a reasonably prudent monetary and fiscal stance, with average inflation at 0.8 percent (see Table 1). In general, after each exit from the Gold Standard, monetary and fiscal policies were aimed at returning within a reasonable time frame to the Standard, albeit at a lower parity. Thus, R-II was relatively high on average (4.6%), and higher than those of the main world financial centers (e.g., the United Kingdom).

From 1914 onwards, there are three major episodes of fiscal deficits: before and during the two World Wars, and during the period of increasing fiscal expansion post-1960. From the end of Bretton Woods in 1971 until the birth of the European Monetary System in 1979 there was also a gap in external exchange rate discipline. This allowed monetary policy to accommodate fiscal policy, and led to increasing inflationary pressures. Post-1979, real interest rates returned to much higher and stable levels (R-II was 5.1% on average), with inflation in the 1980s much lower on average (9.2%) than in the 1970s (14.2%). This resulted from a combination of three factors. First, Italy’s entry into the European exchange rate system in 1979; second, a progressive decoupling of monetary policy from fiscal policy, starting in 1981 through the removal of the obligation on the part of the Bank of Italy to purchase unsold Treasury Bills at auction; and finally, the gradual removal of capital controls, which required a partial convergence with world interest rates.

This brief overview of Italian monetary history shows the changing nature of monetary and fiscal policy over the last 150 years. The econometric results in the next two sections have to be interpreted in the light of the above events.
3. ECONOMETRIC EVIDENCE ON THE FISHER EFFECT AND THE DETERMINATION OF NOMINAL INTEREST RATES

Tests of the Fisher hypothesis (see Fisher, 1930), can be grouped into two broad categories. First, direct estimates of the Fisher equation of the type (Eq. 1):

\[ R_t = \rho + \beta E(\Pi_t|\Omega_{t-1}) + u_t \]  

where test \( \beta = 1 \), where \( \rho \) is the real interest rate, and \( \Omega \) is the information set at time \( t - 1 \). The equation may be estimated either using the errors-in-variables methods, or the two-stage estimation methods where the expected inflation rate is modeled using information available to economic agents at time \( t - 1 \). Studies such as Summers (1983, 1986) fall into this category. One of the problems with such an approach is that normally \( \rho \) is treated as constant, so that the results are conditional on this heroic assumption. The problems are compounded by the fact that, in many cases, we might find that both \( R \) (and implicitly \( \rho \)) are nonstationary over some periods, while \( \Pi \) is integrated of degree zero \([I(0)]\) (see Rose, 1988).

To avoid such problems, a second strand of this literature has focused on whether the inflation expectations implicit in the term structure of interest rates (see MacDonald and Murphy, 1989; Wallace and Warner, 1993) are consistent with the Fisher hypothesis.

Unfortunately in our current study we do not have a data set on the term structure that allows us to take the latter approach, so we have chosen the first approach, based on equation (1) and dynamic variants of this model. We first estimated equation 1 as it stands, using the errors-in-variables approach, with the realized value of current inflation as a regressor, and variables that are in the economic agents’ information set (lagged values of inflation, foreign inflation) as instruments. This has the advantage over the two-stage procedure of producing consistent standard error estimates (see Pesaran, 1987). To examine the evolution of the

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6 On the origins of the Fisher effect see Humphrey (1993).
7 Summers (1983, 1986) uses band-spectrum regressions to verify the hypothesis. This has the advantage of focusing on the long-run movements in the series. We take a similar approach when applying cointegration methods below.
8 Most of the available short-term interest rates were administered rates during much of the pre-World War II period.
Fisher effect over time, we estimated equation 1 as a rolling regression, using a 30 observation window. The estimated coefficient on the inflation variable $\beta$ and its two-times-standard error bands are shown in Figure 2. There are several points to note about these results.

First, for the regression using post-World War II data up to the early 1970s, the coefficient is closer to unity, although always significantly less than 1. This is a result similar to that obtained for the United States by Summers (1986), who found that the 1954–71 period corresponded most closely to the Fisher hypothesis. One reason why the coefficient is likely to fluctuate over time is, of course, the presence of regime shifts, which implies that the hypothesis only holds in the long run (Summers, 1983). Thus, it took some time for inflationary expectations to be built into nominal rates following the higher post-1970 inflation regime, and the Fisher coefficient for the final periods of the sample are consequently lower. This possibility is examined below.

Second, there seems little evidence for the presence of the Fisher effect in the pre-World II period. Closer inspection of the inflation series suggests that inflation in Italy in the interwar period is

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9 In the pre-1914 period, as for the United States and the United Kingdom, it is more difficult to fit an ARMA model for inflation.
forecastable using a restricted ARMA(1,3) model. The inability to capture the Fisher effect before World War II is probably due to a failure of the simplest application of the rational expectations hypothesis. As noted by Barsky and De Long (1991), simple regression models of inflationary expectations and the Fisher hypothesis do not take into account the uncertainty that economic agents face in forming inflationary expectations at a time when inflation does not display any clear trends, and when sudden regime shifts are possible. Overall, it seems clear that expected inflation was equal to zero over the pre-World War II sample.

Third, it may be difficult for such a simple model to discriminate between increases in nominal rates in the 1980s due to the incorporation of inflationary expectations, and positive effects on the real interest rate due to the expansionary fiscal policy in Italy over the last 2 decades.

As a further check on the validity of the Fisher hypothesis post-World War II, we used cointegration methods to estimate the relationship in Equation 1. Cointegration tests and regressions are now commonplace in the applied econometrics literature, and they are aimed at verifying the existence of, and estimating, long-run relationships between integrated time series. (For a full survey of these methods and their development over time see, *inter alia*, Engle and Granger, 1987; Johansen, 1991; Phillips and Loretan, 1991; Banerjee et al., 1993; Muscatelli and Hurn, 1994.) The advantage in using multivariate cointegration approaches compared to a static estimation of Equation 1 is that we explicitly take account of the joint endogeneity of interest rates and inflation (see Phillips and Loretan, 1991; Summers, 1983, 1986; Banerjee et al., 1993).

We focus solely on post-World War II data because that is the only period when inflation displays a clear upward trend [i.e., unit root tests do not reject the null hypothesis that \( \Pi \) is I(1)]. Estimating a four-lag VAR for \( R \) and \( P \) using the procedure proposed by Johansen (1991), we find a single significant cointegrating vector over the period 1948–93 (The maximum eigenvalue statistics are 35.7 and 2.95, and the trace statistics are 38.7 and 2.95, respectively). The estimated coefficient on inflation in the cointegrating vector is 1.138. Testing the null hypothesis of a unit restriction on this coefficient using the likelihood ratio test proposed by

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10 We do not report all the details of this estimation and the graphics of the estimated vectors and eigenvalues for reasons of space. They are available from the authors on request.
Johansen yields a test statistic of 0.77, which is distributed as a \( \chi^2(1) \) under the null, and insignificant at the 5-percent level.

These results suggest that it might have taken some time in the postwar period for inflationary expectations to be fully incorporated into nominal interest rates. But, taking the post-World War II period as a whole, a unit coefficient \( \beta \) can be found, and the Fisher hypothesis is supported by the data.

Having examined the relationship between interest rates and prices, we now turn our attention to quantifying the effect of fiscal factors on real interest rates.

4. REAL INTEREST RATES, FISCAL POLICY, AND NEOCLASSICAL THEORY

In recent years Robert Barro (1987, 1989) and other authors\(^{11}\) have examined the macroeconomic implications of the neoclassical theory of fiscal policy, in contrast to the predictions of the conventional Keynesian aggregative income–expenditure framework. These have been tested, quite successfully, on a long run of U.K. historical interest-rate data by Benjamin and Kochin (1984) and Barro (1987), and for shorter time periods for other developed economies (see Evans, 1987). We now examine whether the predictions of the neoclassical model also apply to Italian interest rates. As noted earlier, the Italian case is of interest because, in comparison to the other G7 countries analyzed in existing studies, it has experienced a much more dramatic public expenditure expansion and debt accumulation in the post-World War II period. We are restricted to estimating our econometric models for the period 1860–1993, due to the lack of a consistent series for fiscal expenditures for the preunification period.

We focus on two particular predictions of the neoclassical model, namely that only temporary increases in government spending increase real interest rates, and that the level of debt accumulated has no impact on the level of real interest rates. These two predictions are tested respectively in Sections 4A and 4B.

4A. The Neoclassical Model

Barro (1987) employs simple least-squares or IV estimates of the following equation to test the neoclassical hypothesis:

\( \text{\textsuperscript{11}} \text{See, for instance, Ahmed (1986) and Aschauer (1985).} \)
\[ R_t - \Pi_t = c + \sum_{j=0}^{\infty} \alpha_j (g - \bar{g})_{t-j} + u_t \]  

(2)

where \( g \) is the ratio of actual government spending to trend real GNP and \( \bar{g} \) is a measure of its “permanent” component. The hypothesis is, therefore, that only transitory increases in spending have an effect on real interest rates. Barro allows for the error term to follow an AR(1) process in his estimated equations, and finds that current and lagged temporary spending effects are important in affecting real interest rates.

In estimating this model we are faced with several issues. First, when estimating equation 2, Barro finds that the AR(1) model he fits for the error term has an estimated coefficient close to unity. The nonstationarity of the residuals is a problem in the U.K. case, as for the period 1701–1918 the U.K. long rate is very volatile, and appears to have a unit root. Thus, Barro’s (1987) regressions are essentially “unbalanced” (see Banerjee et al., 1993; Phillips and Loretan, 1991), in terms of the orders of integration of the dependent and explanatory variables. This is because his only explanatory variable (temporary increases in government military spending) appears to be stationary for most of the sample period, with a few punctuations due to the periods of war.12 In contrast, once one adjusts the Italian long rate for inflation, it is found that, with the exception of the latter part of World War II and the immediate aftermath, the real rate of interest appears to have a relatively constant sample mean and variance (see Figure 1 and Table 1).

The second issue is the definition of the real interest rate. In Barro’s (1987) study of the United Kingdom this is relatively straightforward, and the United Kingdom was almost continuously on the Gold Standard between 1701–1918 (with the exception of the Napoleonic Wars and the First World War), and British wholesale prices tended to follow a random walk. He, therefore, was able to take the nominal interest rate as equivalent to the real interest rate. In the case of Italy, our sample includes post-1918 data, and some allowance has to be made for inflation. As already noted in Section 3, in the interwar years, inflation was predictable, but did not display any general trend. Because of the difficulty in interpreting the pre-World War II data, we experimented with two alternative measures of the real interest rate.

12 See Perron (1989) for an analysis of unit root tests applied to data with structural breaks in them.
The first, RR1 simply sets the real interest rate equal to R-II over the full sample. The second measure (RR2) takes the real interest rate as being simply equal to the nominal interest rate in the pre-1940 period (i.e., we assume that any inflation during these earlier years was unanticipated), and as equal to R-II in the post-1940 period.13

Third, we face the problem of how to construct an appropriate series for temporary changes in government spending. In Barro’s case this was done using a standard “permanent income” definition, where \( g^p \) is defined as a weighted average of future expected government spending. But Barro used military spending as his \( g \) variable, which tended to be relatively constant in peace time, with several sharp peaks at times of war. Thus, in his case, actual and temporary spending tend to follow a very similar pattern. In contrast, we employ the total government spending ratio, and this has tended to display some local trends, with periods of sharp variation. To distinguish between trend and cycle in \( g \), we have employed the Hodrick-Prescott (1980) filter, which defines the trend component at time \( t \), \( g_t^p \) for the sample \( t = 1, \ldots, T \) as the sequence that minimizes Equation 3:

\[
\sum_{t=1}^{T} (g_t - g_t^p)^2 + \lambda \sum_{t=2}^{T-1} [(g^p_{t+1} - g_t^p) - (g^p_t - g_{t-1}^p)]^2
\]

where \( \lambda \) depends on the frequency of the data (see Hodrick and Prescott, 1980). In Figure 3, we plot the difference between \( g \) and the filtered series, \( g^p \), which measures temporary deviations in Italy’s government spending ratio.

The final issue is how one should allow for other influences on the real interest rate in addition to those of temporary government spending. Unlike Barro (1987), this is slightly less critical in our case, as our real interest and temporary spending variables both appear to be integrated of the same order [they are both \( I(0) \)], so that we do not face the problem of nonstationary residuals. However, there is still a potential problem of misspecification bias due to the problem of omitted influences on real interest rates. We tackle this issue in two ways. First, by allowing for an autoregressive element (lags of the dependent variable) in equation 2,

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13We also experimented with more sophisticated expectations-formation mechanisms to construct alternative measures for expected inflation and the real interest rate. For instance, we estimated a structural time series model for inflation (see Harvey, 1989) using the Kalman Filter and used it to construct one-step ahead forecasts for the inflation variable. However, the results obtained do not differ markedly from those reported above.
so that it becomes a more conventional autoregressive distributed lag (ADL) model. Second, as in practice it is quite difficult to identify ARMA models, we try to account for unexplained influences by building a structural time series model for RR with $(g - g')$ as an additional explanatory variable.

The advantage of structural time series models is that they provide a more natural way of decomposing a series into its level, trend, and cyclical components than ARIMA models. Their advantage over ARIMA models lie in the fact that it is simpler to identify the correct specification of structural models, and that simple specifications of structural time series models (i.e., models with a small number of estimated parameters) can be shown to have quite complex ARIMA processes as their reduced form (see Harvey, 1989). A further advantage in our context is that, although our real interest series seems $I(0)$ when tested over the whole sample, across different policy regimes the interest rate might display local trends.

The basic structural time series model that we estimate has the general form:

$$RR_t = \mu + \psi_t + \gamma(g - g') + \epsilon_t$$

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t$$
\[ \beta_t = \beta_{t-1} + \zeta_t \]
\[
\begin{bmatrix}
\psi_t \\
\psi^*_t
\end{bmatrix} = \rho \begin{bmatrix}
\cos(\lambda_c) & \sin(\lambda_c) \\
-\sin(\lambda_c) & \cos(\lambda_c)
\end{bmatrix} \begin{bmatrix}
\psi_{t-1} \\
\psi^*_{t-1}
\end{bmatrix} + \begin{bmatrix}
\kappa_t \\
\kappa^*_t
\end{bmatrix}
\]

(4)

where the model for the interest rate, in addition to the fiscal policy explanatory variable, contains several elements. First a level component, \( \mu \), which can be stochastic, and which can also have a stochastic trend term, \( \beta \). The additional terms are \( \psi \), which is a stochastic cycle, whose evolution is specified in Equation 4. In the cyclical element, \( \rho \) is the damping factor, and \( \lambda_c \) is the frequency measured in radians. This type of full structural time series model has a reduced form that is an ARIMA(2,2,4) process. Because, in our case, the real interest rate seems to be stationary, displaying a trough\(^{14}\) around World War II, we should find that a constant level term with no trend would be a better approximation of the behavior of \( RR \). If the model collapses to a cycle plus trend model with a stationary cycle, then it has a reduced form that is equal to an ARMA(2,2) process.

The structural time series model is estimated using maximum likelihood (ML), over different sample periods, to check whether the effect of temporary government spending is significant. This enables us to check whether the results from the simple ADL regression model are robust to the specification used, given the problem of misspecification bias.

The results from the estimation of the ADL model for real interest rates are reported in Table 2. As the real interest rate behaves very erratically in the World War years due to the sudden jump in inflation, we have included, where this was significant, a step dummy for the World War years (DWAR). The AR(m) statistic is the LM test for serial correlation. The test statistics have to be interpreted with care, however, as the residuals fail both the normality test and the ARCH test when the war years are included. This is not surprising, given the behavior of the real interest rate during these years. For this reason, we also ran the same regressions over subsamples that excluded the war years. We also report the long-run impact of the temporary increase in government spending (LRE) when the equation contains distributed lags or an autoregressive term. For each LRE, we also report the corresponding asymptotic standard error.

\(^{14}\)To take account of the anomalous behavior of the real interest rate in the World War II, we introduce a step dummy for these years in the structural time series model.
### Table 2: ADL Model Estimates for the Real Interest Rate

<table>
<thead>
<tr>
<th>Sample Period</th>
<th>Model Equation</th>
<th>$R^2$</th>
<th>SEE</th>
<th>DW</th>
<th>AR(1)</th>
<th>LRE</th>
</tr>
</thead>
<tbody>
<tr>
<td>1865–1993</td>
<td>$RR_1 = 0.04 - 0.78\text{*DWAR} + 0.70(g-g^p) - 0.64(g-g^p)_{-1}$</td>
<td>0.50</td>
<td>0.23</td>
<td>2.03</td>
<td>0.05</td>
<td>0.06</td>
</tr>
<tr>
<td></td>
<td>(0.02) (0.10)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1948–93</td>
<td>$RR_1 = 0.012 + 0.12(g-g^p) + 0.37\text{*RR}<em>{1-1} - 0.05\text{*RR}</em>{1-4}$</td>
<td>0.43</td>
<td>0.03</td>
<td>1.57</td>
<td>0.19</td>
<td>0.18</td>
</tr>
<tr>
<td></td>
<td>(0.005) (0.03)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1865–1939</td>
<td>$RR_1 = 0.025 - 0.19\text{*DWAR} - 0.06(g-g^p)<em>{-1} + 0.32\text{*RR}</em>{1-1}$</td>
<td>0.57</td>
<td>0.06</td>
<td>2.11</td>
<td>0.79</td>
<td>-0.09</td>
</tr>
<tr>
<td></td>
<td>(0.008) (0.04)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1865–1993</td>
<td>$RR_2 = 0.016 + 0.010(g-g^p) - 0.07(g-g^p)<em>{-2} + 0.35\text{*RR}</em>{2-1} + 0.19\text{*RR}_{2-2}$</td>
<td>0.27</td>
<td>0.05</td>
<td>2.01</td>
<td>0.10</td>
<td>0.07</td>
</tr>
<tr>
<td></td>
<td>(0.006) (0.02)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>$RR_2 = 0.003 + 0.002(g-g^p) + 1.23\text{*RR}<em>{2-1} - 0.29\text{*RR}</em>{2-2}$</td>
<td>0.94</td>
<td>0.004</td>
<td>1.94</td>
<td>0.44</td>
<td>0.03</td>
</tr>
<tr>
<td></td>
<td>(0.006) (0.018)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

**Note:** Standard errors reported in parentheses.

The full-sample results reported in Table 2 are poor. Over the full sample, the fiscal effect on $RR_1$ and $RR_2$ is insignificant (and wrongly signed for $RR_1$). The main problem seems to be the pre-World War II data. Estimating the model over the period up to 1939 yields results that are contradictory to the neoclassical model, with an insignificant coefficient on $(g-g^p)$. But there is strong support for the neoclassical model in the post-World War II years. The total effect of the fiscal variable is significantly positive for the years 1948–93.\(^{15}\)

To provide a further check on the ADL model results, we estimated a structural time series model for $RR_1$ and $RR_2$ to see whether this might help to reconcile some of our contradictory results in Table 2.

We estimated the structural time series model over the same three sample periods. The estimated hyperparameters of the model (the parameters that govern the stochastic movements of the state variables) and the structural parameters are reported in

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\(^{15}\)We start our post-World War II sample in 1948, as the postwar inflation stabilization program was not completed until that year.
Table 3: Structural Time Series Model for the Real Interest Rate

<table>
<thead>
<tr>
<th>Dependent variable: RR1 Sample period 1863–1993</th>
<th>$\sigma^2(\eta) = 0.09$</th>
<th>$\sigma^2(\kappa) = 0.06$</th>
<th>$\sigma^2(\varepsilon) = 0.04$</th>
<th>$\rho = 0.85$</th>
<th>$\lambda_e = 0.74$</th>
<th>$\gamma = 0.41$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(0.11)</td>
<td>(0.06)</td>
<td>(0.01)</td>
<td>(0.12)</td>
<td>(0.12)</td>
<td>(0.13)</td>
</tr>
<tr>
<td>Dependent variable: RR1 Sample period 1948–93</td>
<td>$\sigma^2(\eta) = 0.014$</td>
<td>$\sigma^2(\kappa) = 0.04$</td>
<td>$\sigma^2(\varepsilon) = 2.67$</td>
<td>$\rho = 0.66$</td>
<td>$\lambda_e = 0.000$</td>
<td>$\gamma = 0.08$</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.10)</td>
<td>(0.29)</td>
<td>(0.243)</td>
<td>(0.04)</td>
<td></td>
</tr>
<tr>
<td>Dependent variable: RR2 Sample period 1863–1993</td>
<td>$\sigma^2(\eta) = 0.16$</td>
<td>$\sigma^2(\kappa) = 0.01$</td>
<td>$\sigma^2(\varepsilon) = 0.001$</td>
<td>$\rho = 0.51$</td>
<td>$\lambda_e = 0.56$</td>
<td>$\gamma = 0.01$</td>
</tr>
<tr>
<td></td>
<td>(0.07)</td>
<td>(0.37)</td>
<td>(0.03)</td>
<td></td>
<td>(0.06)</td>
<td></td>
</tr>
<tr>
<td>Dependent variable: RR2 Sample period 1863–1939</td>
<td>$\sigma^2(\eta) = 0.14$</td>
<td>$\sigma^2(\kappa) = 0.01$</td>
<td>$\sigma^2(\varepsilon) = 0.001$</td>
<td>$\rho = 0.97$</td>
<td>$\lambda_e = 0.08$</td>
<td>$\gamma = 0.004$</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.020)</td>
<td>(0.02)</td>
<td>(0.06)</td>
<td>(0.003)</td>
<td></td>
</tr>
</tbody>
</table>

Note: Standard errors reported in parentheses.

Table 3. Note that, with the exception of the full sample estimates (1863–1993) a stochastic level is not necessary [i.e., $\sigma^2(\eta)$], given the stationary nature of both RR1 and RR2 over the subsamples. Similarly, a stochastic trend is not required for any of the models, and hence, estimates for $\sigma^2(\xi)$ are not reported.

The structural time series model yields clearer results for the full sample. For both the RR1 and RR2 models the fiscal variable is significant, in contrast to the ADL model. Under the RR1 definition, even once a stochastic cycle is fitted to the data, the variable is still significant, which suggests that it is not picking up extraneous factors. For the 1948–93 subsample, the structural model essentially confirms the results obtained using the standard ADL model, which support the neoclassical model. For the pre-World War II period, the fiscal variable is not significant, but we do not obtain an anomalous negative estimated coefficient.

The contrasting results between the two different modeling approaches could be explained by the fact that the structural model may be more flexible in capturing nonmodeled influences on the real interest rate through the structural elements of the model. There are three possible explanations for the less significant results obtained for the pre-1939 sample: (i) our two measures of the real interest rate pre-1939 are inadequate, due to the inherent difficulties of measuring inflationary expectations; (ii) there was a major change in the fiscal regime post-1945, and this should have...
been recognized by refining the construction of the permanent component of $g$; (iii) the simple neoclassical model needs to be modified to explain the behavior of interest rates before 1939. For instance, some account should be taken of the effect of the threat of war on assets and hence on real interest rates. These issues could be investigated in further work.

We next turn to the possible role of debt in explaining real interest rates. There seems to be some indication that our models in Table 3 tend to underpredict the real interest rate from the 1980s onwards.

4B. Debt Accumulation and Real Interest Rates

The standard Ricardian neoclassical approach to fiscal policy suggests that debt accumulation should not influence the real interest rate (see Barro, 1987, 1989). Italy has experienced a major debt accumulation since the 1970s, and hence, provides us with an ideal context in which to test this proposition. The ratio has tended to be high throughout the historical period, but has risen sharply in the post-World War II period. We estimate our models over this latter period.

We reestimate both the ADL and structural time series models for the three sample periods, with debt as an additional explanatory variable. Thus, the estimated equations become, for the ADL and structural time series models:

\[ RR_t = c + \sum_{i=0}^{n} \alpha_i (g - g^o)_{-i} + \sum_{i=0}^{n} \beta_i d_{-i} + u_t \]  \hfill (5a)

\[ RR_t = \mu_t + \psi_t + \sum_{i=0}^{n} \gamma_i (g - g^o)_{-i} + \sum_{i=0}^{n} \beta_i d_{-i} + \epsilon_t \]  \hfill (5b)

where $d$ is the debt-to-GNP ratio, and we also experimented with lags of the fiscal variables in Equation 5b to find the best fit.

The results for the sample period 1948–1993 are reported in Table 4. Once again, for the ADL model, LRE$(g)$ and LRE$(d)$ show the long-run effect for each of the two fiscal explanatory variables. For the ADL model, the debt variable turns out to have a significant positive long-run effect in the post-World War II years, in addition to the government spending variable. This is confirmed using the structural time series model, where, once again the total effect of the $d$ variable turns out to be positive and significant. These results are in sharp contrast to the predictions of the debt equivalence models, and suggest that Italian real interest
Table 4: The Effect of Debt on Real Interest Rates

<table>
<thead>
<tr>
<th>Model</th>
<th>Dependent variable</th>
<th>Sample period</th>
<th>$RR_{1}$</th>
<th>$g_{t-1}$</th>
<th>$g_{t-2}$</th>
<th>$d_{t-1}$</th>
<th>$d_{t-2}$</th>
<th>$d_{t-4}$</th>
<th>$dt_{21}$</th>
<th>$R^{2}$</th>
<th>$SEE$</th>
<th>$DW$</th>
<th>$AR(1)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>ADL model</td>
<td>$RR_{1}$</td>
<td>1948-93</td>
<td>$-0.001 + 0.08\times (g_{t-1}) + 0.34\times RR_{t-1} + 0.06\times RR_{t-4} + 0.04\times d_{t-1}$</td>
<td>$(0.008)$</td>
<td>$(0.10)$</td>
<td>$(0.02)$</td>
<td>$(0.016)$</td>
<td>$(0.008)$</td>
<td>$0.47$</td>
<td>$0.039$</td>
<td>$1.83$</td>
<td>$0.52$</td>
<td></td>
</tr>
<tr>
<td>Structural time series model</td>
<td>$RR_{1}$</td>
<td>1948-93</td>
<td>$\sigma^{2}(\eta) = 0.02 \quad \sigma^{2}(\epsilon) = 0.04 \quad \rho = 0.62 \quad \lambda_1 = 0.00 \quad \gamma_0 = 0.12$</td>
<td>$(0.01)$</td>
<td>$(0.01)$</td>
<td>$(0.28)$</td>
<td>$(0.25)$</td>
<td>$(0.04)$</td>
<td>$0.62$</td>
<td>$0.33$</td>
<td>$0.39$</td>
<td>$(0.13)$</td>
<td>$(0.14)$</td>
</tr>
</tbody>
</table>

rates are unlikely to fall much without some effort to reduce the debt ratio. They tend to validate alternative theories of fiscal policies that stress the impact of the stock of public debt on interest rates (see Blanchard, 1985).

Finally, further investigation of the ADL regression in Table 4 using rolling regression methods showed that the debt variable gains significance around 1980, which corresponds to the point at which monetary policy began to be less accommodating. This is shown in Figure 4, which plots the estimated coefficient and standard error bands on the $d_{t-1}$ regressor using a 15-observation window rolling regression. There might, therefore, be some role for the method of financing of deficits and the refinancing of debt stocks in the determination of real interest rates. The evidence might also suggest that the effect of the debt stock on real interest rates only becomes detectable at some threshold value.

5. CONCLUSIONS

In this paper we have examined some of the factors underlying the determination of nominal and real interest rates in Italy over a long historical period. There are two main conclusions from our study.

First, our evidence tends to show, in line with similar evidence for the US (see Summers, 1983, 1986; Wallace and Warner, 1993), that the adjustment of nominal rates to expected inflation tended to be rather imperfect and slow, even in the post-World War II period. There is also evidence that in the pre-1939 period, the
effect was totally absent. Nevertheless, taking the 1948–90 period as a whole, cointegration methods suggest that a unit coefficient restriction on the inflation variable cannot be rejected, validating the Fisher hypothesis.\textsuperscript{16} Overall, this might suggest that inflationary expectations only adjust following sustained changes in the local trend of inflation, and that expectations do not adjust in response to short-run movements in inflation that are predictable. Barsky and DeLong (1991) provide an attractive explanation for this phenomenon, which might be attributable to uncertainty in the underlying economic structure.

Second, there seems to be evidence contrary to the pure neoclassical model of fiscal policy in post-World War II Italy. Although there is evidence that deviations of the government spending ratio from trend have a positive impact on real interest rates, particularly post-World War II, the increasing debt ratio also seems to be exerting a positive influence on real interest rates, along the lines suggested in Blanchard (1985).

\textsuperscript{16}All this seems to confirm Schwartz’s statement that “... There (now) is less short-term but more long-term variability in rates of inflation and much higher levels of inflation than had been experienced in peace time over the past century. As a result, market participants have a greater incentive to seek to allow for future price movements ...” Schwartz (1987).
REFERENCES


