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A FUNCTIONAL COEFFICIENT APPROACH TO MODELING THE FISHER HYPOTHESIS: WORLDWIDE EVIDENCE

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We pursue a semiparametric approach to examining core implications of the Fisher hypothesis, namely cointegration linking nominal interest rates and inflation, and homogeneity of the potential equilibrium relation. The sample is an unbalanced panel and comprises monthly time series from more than 100 economies. The time period of at most 45 years is subdivided into three regimes according to dominating monetary policies. To exploit the cross-sectional dimension for inference on parameter homogeneity, we apply mean group estimation of functional coefficients that allow the conditioning of key model parameters on economic states. The evidence in favor of cointegration is weakened over states of negative real interest rates that are likely to coincide with scenarios of high inflation. The ex post real interest rate is mostly diagnosed as unstable. The Fisher hypothesis is particularly confirmed for states characterized by large positive interest rate adjustments during the inflation-targeting regime.

Keywords: Fisher Hypothesis, Panel Data, Functional Coefficient Models

1. INTRODUCTION

The relationship between nominal interest rates and inflation has been investigated frequently in both theoretical and empirical economics. Fisher (1930) formalized a model where nominal interest rates respond one to one to expected changes of the price level. This is typically referred to as the Fisher hypothesis. The Fisher coefficient may differ from unity when interest income is subjected to taxation [Crowder and Hoffman (1996)]. Moreover, in the case of accelerating inflation, the basic link between interest rates and inflation might be disturbed by agents shifting out of nominal and acquiring real assets [Tobin (1965)]. Recent contributions to modeling monetary policy via policy rules [Taylor (1993), Woodford (2003)] could also imply that the Fisher coefficient differs from unity.

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Empirical studies for the United States by Fama (1975) and Fama and Schwert (1977) document evidence in favor of approximately constant real interest rates, as implied by the Fisher hypothesis. Beginning in the 1990s, the Fisher hypothesis underwent empirical tests that take the potential nonstationarity and cointegration of the involved time series explicitly into account [Mishkin (1992)]. Finding evidence in favor of a cointegration relationship for the United States, Evans and Lewis (1995) document a long-run coefficient less than unity, whereas a coefficient estimate in excess of unity is found by Crowder and Hoffman (1996).

In the analysis of a set of macroeconomies, pooling is a promising device to improve the efficiency of statistical procedures. Panel cointegration methods have been attracting a great interest in the recent econometric literature [Banerjee (1999)]. Phillips and Moon (1999) emphasize that for panel data models, challenges posed by the statistical analysis of nonstationary processes, such as spurious regression, may be addressed by exploiting the variation of parameter estimates over the cross section. From this perspective, the adoption of panel data econometrics might be particularly fruitful if economic theory implies specific and unique parametric restrictions such as the postulate of purchasing power parity. If, however, economic theory is in line with a range of admissible parameter values, the a priori merits of a panel approach are less clear. With respect to the size of the Fisher coefficient, economic theory is not conclusive, so that panel heterogeneity is likely.

Apart from cross-sectional heterogeneity the relation between interest and inflation rates could also be driven by economic factors changing over time, such as tax regulations or regimes of monetary policy. With respect to empirical modeling, this argument applies in particular if longitudinal data are investigated. Econometric models, error correction specifications say, typically proceed from the assumption that model parameters are invariant over time. Parameter invariance might be regarded as a strong restriction. Consequently, recent contributions to the econometric analysis of cointegrated systems allow for nonlinearity of the equilibrium relationship or adjustment toward the latter, e.g., Balke and Fomby (1997), Granger (2001), Park and Phillips (2001), Chang et al. (2001), Escribano (2004), and Karlsen et al. (2007).

In this paper, we investigate the relation between nominal interest and inflation rates for an unbalanced cross section including the majority of the world's economies. Monthly time series covering a period of at most 45 years (1960–2004) enter the analysis. Given the likelihood of cross-sectional heterogeneity, nonlinearity, and parameter variation over time, we rely on country-specific regressions with parameters that are allowed to depend on economic states. Owing to the large time dimension, it is possible, if not likely, that the dynamic features of the real interest rate correspond to exchange rate regimes, changing monetary policy rules, tax regulations, etc. Notably, time windows over which the latter conditions change are eventually too large to capture dynamic real interest rate features by means of smooth transition error–correction models. For the latter reason, we adopt a functional coefficient approach introduced by Cai et al. (2000).

In this model class, standard specifications of cointegrating variables are treated in a semiparametric fashion by conditioning the model parameters on measurable economic states. State dependence implies nonlinear dynamics between interest rates and inflation and may also be seen as a general framework for testing against time-invariant models. Moreover, unconditional full-sample analyses are complemented by time-specific modeling where three distinct subperiods (1960–1978, 1979–1989, and 1990–2004) are investigated.

To preview some results, for the majority of countries interest and inflation rates are diagnosed as jointly nonstationary. The evidence in favor of cointegration is weakened over states of negative real interest rates, which are likely to coincide with scenarios of high inflation. Even in cointegrated systems, the ex post real interest rate is mostly unstable. The Fisher hypothesis is particularly confirmed for states characterized by large positive interest rate adjustments during the inflation-targeting regime. For this subperiod, the mean group Fisher coefficient exceeds unity conditional on states of small nominal interest rates. Moreover, the inflation-targeting period is characterized by error-correction dynamics operating locally through the adjustment of inflation. As a force contributing to price stability, the inflation expectation channel is known to require high credibility of the monetary authorities.

The remainder of this paper is organized as follows: In the next section we briefly sketch the economic model and its econometric counterpart. The data and a few estimation results obtained from standard linear models are discussed. Implementation and specification issues arising in the framework of functional coefficient models are in the focus of Section 3. Empirical results obtained from the latter approach are provided in Section 4. Section 5 summarizes the main findings and concludes.

2. PARAMETRIC MODELING OF THE FISHER RELATION

2.1. The Economic Model

According to the Fisher equation the nominal interest rate in time t (R_t) is composed of the ex-ante real interest rate ($E_{t-1}[r_t]$) and the ex-ante expected inflation rate ($E_{t-1}[\pi_t]$) [Mishkin (2003)]. Formally one has

$$R_t = E_{t-1}[r_t] + E_{t-1}[\pi_t] + u_t, \tag{1}$$

where $E_{t-1}[\bullet]$ denotes the expectations operator conditioning on information available in time $t-1$ and u_t is stationary zero-mean disturbance term. Under rational expectations the expected and the actual inflation rate differ by a stationary zero-mean forecast error. The inflation and nominal interest rate are observable. Thus, the ex post real interest rate is

$$r_t = R_t - \pi_t + v_t, \tag{2}$$

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where υ_t is a stationary zero-mean error term. Equation (2) provides a basis for the econometric strategy to test the Fisher hypothesis. Assuming υ_t to be stationary, the integration properties of r_t are determined by the corresponding features of R_t and π_t. If the latter variables are both stationary $[R_t,\ \pi_t \sim I(0)]$, then $r_t \sim I(0)$. In case one variable is nonstationary and the other variable is stationary, the real rate is nonstationary $[r_t \sim I(1)]$, which is at odds with the Fisher hypothesis. If both variables are nonstationary $[R_t,\ \pi_t \sim I(1)]$ and the linear combination $R_t - \pi_t$ is stationary the Fisher hypothesis implies a $(1,\ -1)'$ cointegrating vector linking interest rates and inflation.

For these considerations the empirical analyses in this work are conditional on a separation of the cross section according to diagnosed stochastic trending features. The particular subsamples comprise cross-section members for which nominal interest and inflation rates are both diagnosed as either stationary or nonstationary. A third group of economies is characterized by distinct degrees of integration of both variables. Because this particular group features weakest correspondence to the Fisher hypothesis, "deeper" empirical exercises concentrate on cross-section members for which the involved time series share the same order of integration.

Numerous contributions have been put forth to explain cointegration between nominal interest rates and inflation with an unrestricted cointegration parameter. Accounting for tax rates could imply a Fisher coefficient that differs from unity [Crowder and Hoffman (1996)]. Moreover, Mishkin (1992) argues against an one-to-one link between inflation and interest rates under prevalence of risk premia. According to portfolio theory, risk-averse investors deserve compensations against inflation risk, because unanticipated inflation diminishes the real return on investments. D'Amico et al. (2008) document time-varying inflation risk premia for the United States that contribute to the yield spread of nominal and inflation-protected treasury securities. Furthermore, Söderlind (2008) shows that inflation uncertainty is an important factor explaining risk premia in the United States. Finally, recent approaches to model the monetary policy of central banks by means of Taylor-type policy rules overcome the restrictive presumption of a one-to-one relationship [Taylor (1993), Woodford (2003)].

2.2. The Econometric Model

To investigate the long-run Fisher relation, consider a country-specific bivariate vector error correction model (VECM)

$$\Delta \mathbf{y}_{it} = \nu_i + F_i \mathbf{y}_{it-1} + \Gamma_{i1}\Delta \mathbf{y}_{it-1} + \cdots + \Gamma_{ip}\Delta \mathbf{y}_{it-p} + e_{it}, \quad t = 1, \ldots, T_i, \qquad \mathbf{(3)}$$

where $\mathbf{y}_{it} = (R_{it}, \pi_{it})'$ collects observations of short-term nominal interest and inflation rates for time t in country i, $i = 1, \ldots, N$. Note that the vector of intercept parameters ν_i, the (2×2)-dimensional parameter matrices governing short-run dynamics $\Gamma_{ik}, k = 1, \ldots, p$, and the matrix F_i are cross section–dependent.

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For convenience of model representation and implementation, we assume that presample values are available and that the autoregressive order p is unique over the cross section. In the empirical analysis of monthly data $p = 2$ is chosen throughout. By assumption, e_{it} is a serially uncorrelated residual vector with mean zero and covariance matrix Ω_i. If the variables in $\mathbf{y}_{it}$ are integrated of order 1 and cointegrated with cointegration rank $r = 1$, the matrix F_i allows a factorization $F_i = a_i b_i'$, where both a_i and b_i are 2×1 vectors.

For the purpose of straightforward model implementation and estimation, we focus on single-equation models that can be derived from the VECM in (3). Because our interest is in the cross-sectional pattern of parameter estimates, we regard the efficiency loss involved with country-specific single-equation analyses as negligible. Interest rate and inflation adjustments are formalized, respectively, as

$$\Delta R_{it} = \nu_{1i} + \alpha_{1i}(R_{it-1} - \theta_i \pi_{it-1}) + \sum_{k=1}^{2} \psi_{ik}^{(1)} \Delta \pi_{it-k} + \sum_{k=1}^{2} \phi_{ik}^{(1)} \Delta R_{it-k} + e_{it}^{(1)}, \tag{4}$$

$$\Delta \pi_{it} = \nu_{2i} + \alpha_{2i}(R_{it-1} - \theta_i \pi_{it-1}) + \sum_{k=1}^{2} \psi_{ik}^{(2)} \Delta \pi_{it-k} + \sum_{k=1}^{2} \phi_{ik}^{(2)} \Delta R_{it-k} + e_{it}^{(2)}. \tag{5}$$

Because an ECM estimate of the cointegration parameter in (4) [or (5)] is the ratio of two OLS coefficient estimates, $-\hat{\theta}_i = \widehat{\theta_i \alpha_{1i}}/\hat{\alpha}_{1i}$ (or $-\hat{\theta}_i = \widehat{\theta_i \alpha_{2i}}/\hat{\alpha}_{2i}$), one may expect a priori a few outlying estimates over the cross section. Therefore, we rely on (4) [or (5)] to address the issues of cointegration and weak exogeneity and estimate the Fisher coefficient θ_i by means of dynamic OLS (DOLS) regression models [Saikkonen (1991), Stock and Watson (1993)],

$$R_{it} = \mu_i + \theta_i \pi_{it} + \sum_{k=-2}^{2\backslash 0} \delta_{ik} \Delta \pi_{it+k} + u_{it}, \tag{6}$$

where summation up to "$2 \backslash 0$" indicates that quantities indexed with $k = 0$ do not enter the sum. Error terms u_{it} in (6) are assumed Gaussian and, by construction, independent from $\Delta \pi_{it+k}$, $k = -2, -1, 1, 2$. The model in (2) implies that $\theta_i = 1$ in (4), (5), and (6).

2.3. Data

Data are taken from the international financial statistics of the International Monetary Fund (IMF). Time series are sampled at the monthly frequency and span at most the period from 1960:1 to 2004:6. The unbalanced panel comprises time series from 114 economies. A list of the considered countries is given in the

Appendix. Counting over both data dimensions, almost 32,000 observations enter the empirical analyses. The annual inflation rate is determined from national consumer-price indices (P_{it}) as $\pi_{it} = 100(\ln P_{it} - \ln P_{it-12})$. To measure short-term nominal interest rates, we mostly use money market rates. If these are not available for a member of the cross section, we draw either the treasury bill or the discount rate. Missing values in a series are replaced by linearly interpolated data.

Overall, the sample period of more than four decades covers distinct international regimes and paradigms of monetary policy. It starts in the Bretton Woods era. As a fixed exchange-rate regime, the Bretton Woods system hindered all participating central banks except the Federal Reserve System of the United States (Fed) from conducting an independent monetary policy. The Fed targeted money market conditions [Mishkin (2003, Chapters 18 and 20)]. By coincidence, floating exchange rates were established in the beginning of the 1970s and the Fed shifted its policy focus to the targeting of monetary aggregates. As a consequence of oil price shocks in the 1970s, inflation rates accelerated worldwide, resulting in an obvious need for a strictly disinflationary policy at the end of the 1970s. Owing to factual instability of money demand functions, the Fed weakened its focus on monetary aggregates in the 1980s and early 1990s. In the light of persistently high inflation rates and unstable money demand relations, the monetary authorities in New Zealand (1990), Canada (1991), and the United Kingdom (1992) decided to establish inflation targeting as a new monetary strategy. In the sequel, inflation targeting became an important and often successful strategy of monetary policy worldwide. As a consequence, international inflation rates decelerated. In addition, it is argued that during the last two decades globalization and the failure of the central planning system have been reducing or stabilizing inflation at least for major industrial economies such as the United States, Canada, Western Europe, Japan, and Australia [IMF (2006), Borio and Filardo (2007)].

Given distinct regimes of monetary policy, the unconditional empirical analyses in this study are complemented by conditioning on three subsamples covering the time periods 1960–1978, 1979–1989, and 1990–2004. Providing further support for this separation, Ciccarelli and Mojon (2005) show for 22 OECD members that the first (third) subperiod is characterized by globally increasing (decreasing) inflation rates.

2.4. Some Preliminary Results

Owing to the large cross-sectional dimension, a provision of detailed results on testing for integration or cointegration features of nominal interest or inflation rates is rather space-consuming. Moreover, the focus of this paper is on potential state dependence of core implications of the Fisher hypothesis. Therefore, we interpret only a few aggregated results offered from common linear specifications in the following.

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Diagnostics and subsampling. EC parameters α_{ji} reflect the strength of the interest ($j = 1$) [inflation ($j = 2$)] rate's response to deviations from the long-run equilibrium. Note that in the case of cointegration, at most one of the two EC coefficients might be zero, owing to weak exogeneity of either interest ($\alpha_{1i} = 0$) or inflation rates ($\alpha_{2i} = 0$). For the case where nominal interest and inflation rates are stationary, single-equation ECMs can be reparametrized in terms of an autoregressive distributed lag model such that standard OLS diagnosis applies [Banerjee et al. (1993, Chapter 2)]. In light of these considerations, one may look at the fraction of EC coefficient estimates $\hat{\alpha}_{1i}$ having t-ratios smaller than -1.77, the expected value under the hypothesis of no cointegration [Westerlund (2005)], or smaller than the corresponding 5% critical value (-3.37). In a scenario of cointegration the frequency of t-ratios of $\hat{\alpha}_{1i}$ less than -1.64 could be used to illustrate the significance of EC dynamics of interest rates at a cross-sectional level. Similarly, the frequency of t-ratios of $\hat{\alpha}_{2i}$ exceeding 1.96 in absolute value is a descriptive means to assess weak exogeneity of inflation.

Table 1 summarizes common (parametric) estimates obtained from the models (4) to (6). Apart from full-sample estimates, subsample-specific results are also documented. Owing to unbalancedness of the full cross section, each subsample comprises a distinct number of economies. Subsample data for the full cross section of 114 economies are only available for the period 1990–2004, whereas 42 and 81 countries enter the analyses for the periods 1960–1978 and 1979–1989, respectively. In addition, Table 1 displays separate mean group or quantile statistics for sets of economies that are characterized by distinct time series features of interest rates and inflation. We distinguish economies where both time series are diagnosed to be integrated either of order one ["$I(1)$"] or zero ["$I(0)$"] by means of ADF regressions. The significance level adopted for this classification is 5%. As a third group we list estimation results for economies where one series is diagnosed as nonstationary whereas the other is found stationary ["$I(0, 1)$"]. The detection of distinct degrees of integration is regarded as evidence against Fisher's postulate. For the full sample 44 (of 114) economies are characterized by distinct degrees of integration featuring inflation and interest rates. Such evidence against the Fisher hypothesis is weakest for the first subsample, comprising only 5 (of 42) cross-section members with "odd" degrees of integration. As a reflection of time-varying stochastic features, it is noteworthy that the fraction of economies characterized by joint degrees of integration is higher for all subperiods (between 87/114 and 37/42), as it is for the overall sample (70/114).

For single-country regressions, it turned out that (for particular subsamples) a few estimates of either the EC or cointegrating parameters unreliably exceed an a priori threshold of 10 (!) in absolute value. Single countries characterized by such extreme parameter estimates are removed from the cross section and the number of "outliers" is documented in Table 1. For instance, on a priori reasoning 3 of 35 economies are removed from the "$I(1)$" set of economies and sample period 1960–1978.

TABLE 1. Descriptive statistics for estimated Fisher coefficients and EC parameters

	1960–2004			1960–1978		1979–1989		1990–2004	
	$I(1)$	$I(0)$	$I(0,1)$	$I(1)$	$I(0)$	$I(1)$	$I(0)$	$I(1)$	$I(0)$
					$\hat{\theta}_i$: DOLS				
# i	57	13	44	35	2	60	3	80	7
# $i\ t$	16069	3543	12211	4277	426	6727	372	12293	931
# $\lvert\hat{\theta}_i \vee \hat{\alpha}_{ji}\rvert > 10$	0	0	2	3	0	2	2	1	0
					$\hat{\alpha}_1$: ECM				
mg	−0.027	−0.144	−0.074	−0.100	−0.105	−0.064	−0.182	−0.043	−0.286
	(−4.92)	(−3.35)	(−3.67)	(−3.90)	(−6.03)	(−5.09)	(1.00)	(−5.03)	(−2.71)
min	−0.262	−0.581	−0.625	−0.458	−0.122	−0.499	−0.182	−0.564	−0.773
max	0.006	0.051	0.014	0.083	−0.087	0.138	−0.182	0.013	−0.034
$P(t_{\alpha_1} < -1.64)$	0.614	0.923	0.619	0.594	1.000	0.569	1.000	0.456	0.857
$P(t_{\alpha_1} < -1.77)$	0.561	0.923	0.595	0.500	1.000	0.534	1.000	0.392	0.857
$P(t_{\alpha_1} < -3.34)$	0.070	0.538	0.262	0.031	0.500	0.086	0.000	0.076	0.429
					$\hat{\alpha}_2$: ECM				
mg	0.016	0.014	0.018	0.169	0.002	−0.147	−0.139	0.030	0.017
	(1.76)	(1.52)	(2.00)	(1.07)	(0.07)	(−1.99)	(1.00)	(3.78)	(1.19)
min	−0.133	−0.033	−0.257	−2.297	−0.021	−3.246	−0.139	−0.133	−0.033
max	0.282	0.087	0.157	3.765	0.025	0.559	−0.139	0.282	0.074
$P(\lvert t_{\alpha_2}\rvert > 1.96)$	0.158	0.231	0.238	0.250	0.000	0.190	1.000	0.241	0.429
					$\hat{\theta}_i$: DOLS				
$\lvert\hat{\theta}_i \notin [-1,3]\rvert$	1	1	1	1	0	0	0	0	1
mg	0.437	0.885	0.448	0.270	0.144	0.307	−0.252	0.713	0.699
	(.067)	(.225)	(.095)	(.055)	(.042)	(.060)	(.000)	(.092)	(.075)
					$\hat{\theta}_i$: static regression				
# $\lvert\hat{\theta}_i \notin [-1,3]\rvert$	1	1	1	1	0	0	0	0	1
mg	0.414	0.929	0.418	0.254	0.125	0.291	−0.183	0.668	0.803
	(.062)	(.222)	(.087)	(.049)	(.025)	(.058)	(.000)	(.086)	(.166)

Notes: Columns "$I(0)$," and "$I(1)$," and "$I(0,1)$" correspond to cross-section members for which both series are diagnosed stationary or nonstationary, or interest and inflation rates are characterized by distinct degrees of integration, respectively. Country-specific lag orders in ADF regressions are determined by means of the AIC criterion. Results are given for the entire (unbalanced) panel and for subperiods 1960–1978, 1979–1989, and 1990–2004. The number of countries entering the particular cross sections is given in row "# i." Similarly, "# $i\ t$" provides the number of observations entering the DOLS regression (6). Model-specific observation counts for the regressions (4) or (5) are qualitatively identical. "mg" are mean group estimates from modeling the Fisher relation. Standard error estimates (t-ratios) are given in parentheses underneath mean group estimators of the Fisher coefficient (EC parameters). Cross-sectional minimum and maximum estimates are provided for both EC coefficients. Fisher coefficient estimates are documented for the DOLS and static regression. Mean group estimates of θ_i are conditional upon single estimates $-1 < \theta_i < 3$. The number of removed estimates is given in row "# $\lvert\hat{\theta}_i \notin [-1,3]\rvert$." The remaining entries are frequencies of t-ratios of EC coefficient estimates ($\hat{\alpha}_{ji}$) smaller than -3.37 (5% critical value for testing the hypothesis of no cointegration), -1.77 (expectation of $t_{\hat{\alpha}_1}$ under the hypothesis of no cointegration), and -1.64 (one-sided significance test under cointegration). Weak exogeneity of inflation is assessed in terms of the fraction of t-ratios of $\hat{\alpha}_{2i}$ that exceed 1.96 in absolute value.

Estimating the long-run parameter θ_i, we obtain for a few cross-sectional entities rather large and outlying estimates that are likely to deteriorate the reliability of mean group estimation and inference. Mean group estimates of the Fisher coefficient are provided conditional on single country-specific estimates $-1 \leq \hat{\theta}_i < 3$. The number of outlying coefficient estimates is also given in Table 1. For the full sample, for instance, we count two outlying DOLS estimates, $\hat{\theta}_i \notin [-1; 3]$.

A parametric view of the Fisher relation. In case inflation and nominal interest rates are diagnosed nonstationary ["$I(1)$"], mean group inference for the full sample is uniformly supportive for stabilizing EC dynamics of interest rates that are significant at the 5% level. To better assess the reliability of mean group estimates, Table 1 also documents the empirical range of EC parameter estimates. Estimated EC coefficients for interest rate adjustments ($\hat{\alpha}_{1i}$) vary between −0.773 and 0.138, whereas for the adjustment of inflation, eventually outlying parameter estimates $\hat{\alpha}_{2i}$ (e.g., −3.25, −2.30, 3.77) are found. The adjustment of inflation is mostly insignificantly positive at the mean group level. For the ["$I(1)$"] subsample comprising 58 economies over the time period 1979–1989, the mean group estimate is significantly negative, which is at odds with stabilizing VECM dynamics. This particular result, however, is due to a single outlying (and unreliable) estimate $\hat{\alpha}_{2i} = -3.25$. As a further criterion underpinning weak exogeneity of inflation, it is noteworthy that markedly fewer t-ratios of $\hat{\alpha}_{2i}$ exceed 1.96 in absolute value when t-ratios of $\hat{\alpha}_{1i}$ are below −1.64 or −1.77. For instance, regarding the full sample period for the "$I(1)$" cross section, the hypothesis of weak exogeneity of inflation is rejected for 15.8% of all cross-section members, whereas 56.1% show a t-ratio of $\hat{\alpha}_{1i}$ smaller than −1.77. Comparing the cross sections "$I(0)$" and "$I(1)$," the downward response of interest rates to lagged equilibrium errors is uniformly stronger for the former sample. However, markedly fewer economies are collected in "$I(0)$," so that we do not put too much weight on this result.

Conditional on 57 (13) economies composing the full sample "$I(1)$" ["$I(0)$"] cross section, the mean group Fisher coefficient is 0.437 (0.885). Corresponding standard errors indicate that mean group estimates differ significantly from unity in nonstationary systems. In case both series are diagnosed stationary, the empirical evidence is more in line with Fisher's homogeneity postulate. For the nonstationary systems, it is worth mentioning that, globally, the real interest rate is nonstationary. According to Amato (2005), stochastic trending of real interest rates might reflect imperfections of capital and credit markets, such as risk premia or credit rationing. Trending real rates hint at the risk of monetary strategies implemented under the belief of real-rate stability.

With regard to particular subsamples, it becomes obvious that the Fisher relation hardly holds uniformly over time. The mean group Fisher coefficient is around twice as large for the period of inflation targeting (1990–2004) as for earlier time spans. This shift is consistent with Rogoff (2003) diagnosing a worldwide disinflation trend from 1990 until 2003. After experiencing relatively high inflation in the 1970s, most central banks in industrial economies aimed to stabilize inflation rates at low levels [Rogoff (2003), Romer (2005)]. The strengthened focus of the central bank on the inflation target is reflected in a stronger response of interest rates to inflation. According to Rogoff (2003), the increasing focus of central banks on the inflation target was one reason for the global disinflation at the beginning of the new millennium. Romer (2005) argues that inflation persisted during the 1960s and 1970s because policymakers relied on flawed economic models. According to her, aggregate demand contraction was believed to be hardly effective for inflation control, and monetary authorities refused to adopt anti-inflationary monetary policies. At the end of the 1970s, monetary authorities recognized that inflation was costly and could be controlled by an aggregate demand policy. In the 1990s, some industrial economies fought successfully against high inflation rates by adopting inflation targeting. Opposite to this experience, Central and Eastern European economies faced difficulties in stabilizing their inflation rates at the beginning of the transition process. Taylor (2005) documents that the objective of low (moderate) and stable inflation rates has been strengthened in transition countries and emerging markets in the mid-1990s, say.

In case for a particular economy nominal interest and inflation rates are diagnosed as stationary, the DOLS regression may suffer from estimation inefficiency, as opposed to a simple static regression of interest rates on a constant and the level of inflation. It turns out that for the "$I(0)$" subsample, static and DOLS regressions yield very similar first- and second-order mean group estimates. This impression also holds for the "$I(1)$" cross section. The fact that DOLS and static regression results for the Fisher coefficient are quantitatively very similar holds for all empirical analyses performed in this study including the functional (semiparametric) exercises. Therefore, further results from static regressions are not included, to economize on space.

Overall, preliminary diagnostic results are in line with the established view that interest rates are mostly integrated of order one, whereas inflation rates are diagnosed as stationary for some fraction of the cross section. Similarly to Rose (1988), Koustas and Serletis (1999), Crowder (2003), Driffill and Snell (2003), Rapach and Weber (2004), and Westerlund (2005), our results for the "$I(1)$" cross section are more in favor of time-varying cointegration than against it. Detailed full-sample parametric estimation results for the investigated panel, including a discussion of weak exogeneity, cross-sectional model stability and panel-based approaches to unit root and cointegration testing are provided in Herwartz and Reimers (2006). The latter study also sketches empirical results for the so-called reverted Fisher relation, where inflation is employed as the left hand–side variable in (6). Potential state dependence of EC dynamics and the Fisher coefficient may contribute to biased unconditional estimates. State dependence of the parameters of interest is discussed in Section 4.

3. A FUNCTIONAL COEFFICIENT APPROACH

In analyzing a large cross section of longitudinal data, parameter heterogeneity over both data dimensions is likely. Owing to the large time dimension, economic states impacting on the relation between interest rates and inflation might lack stability for a given economy. For these reasons, the models in (4) to (6) are generalized toward semiparametric specifications with functional coefficients [Cai et al. (2000)]. Functional coefficient models allow conditioning of the Fisher relation on measurable states that could last over long time windows. Because the reader may not be familiar with this model class, we briefly comment below on model representation, bandwidth selection, and semiparametric estimation in turn. Empirical results obtained from the semiparametric models are discussed in Section 4.

3.1. Specification and Estimation

Apart from allowing cross-sectional heterogeneity, we regard the key parameters of the econometric models in (4) to (6), namely θ_i and α_{ji}, $j = 1, 2$, to be time-varying, in the sense that they are dependent on measurable economic states or factors. For notational convenience we refer to any factor as ω and clarify its explicit choice when discussing estimation results.

To concentrate on time variation of the parameters of interest, α_{1i} in (4), α_{2i} in (5), and θ_i in (6), we presume that parameters attached to deterministic or stationary explanatory variables are time-invariant. For this reason, we first apply partial regression techniques to isolate the relation of interest. Then, in a second step, this is generalized toward a functional coefficient model. For each member of the cross-section, model (4) reads compactly as

$$\Delta R_i = (R_{i-}\,, \pi_{i-})\gamma_i + (\mathbf{1}, \Delta R_{i,-1}, \Delta R_{i,-2}, \Delta\pi_{i,-1}, \Delta\pi_{i,-2})\vartheta_i + e_i, \qquad \mathbf{(7)}$$

where R_{i-}, π_{i-} ($\Delta R_{i,\bullet}$, $\Delta\pi_{i,\bullet}$) are T_i-dimensional column vectors collecting the lagged level data R_{it-1}, π_{it-1} (differenced variables ΔR_{it}, $\Delta\pi_{it}$ at the indicated lag) and $\mathbf{1}$ is a unit vector. Accordingly, $\gamma_i = (\alpha_{1i}, \alpha_{1i}\theta_i)'$ and $\vartheta_i = (\nu_{1i}, \psi_{i1}^{(1)}, \psi_{i2}^{(1)}, \phi_{i1}^{(1)}, \phi_{i2}^{(1)})'$. Model (5) obeys an analogous representation (explaining $\Delta\pi_i$ instead of ΔR_i, $\gamma_i = (\alpha_{2i}, \alpha_{2i}\theta_i)'$) and (6) is compactly

$$R_i = \pi_i\gamma_i + (\mathbf{1}, \Delta\pi_{i,-2}, \Delta\pi_{i,-1}, \Delta\pi_{i,+1}, \Delta\pi_{i,+2})\vartheta_i + e_i, \qquad \mathbf{(8)}$$

with $\gamma_i = \theta_i$ and $\vartheta_i = (\mu_i, \delta_{i,-2}, \delta_{i,-1}, \delta_{i1}, \delta_{i2})'$. Both representations (7) and (8) conform with the general notation

$$y_i = Z_i\gamma_i + X_i\vartheta_i + e_i, \qquad \mathbf{(9)}$$

such that γ_i contains the parameter(s) of interest and ϑ_i collects coefficients that are assumed time-invariant. Partialling out X_i from (9) obtains

$$\tilde{y}_{it} = \tilde{z}_{it}'\gamma_i + \tilde{e}_{it}, \qquad \mathbf{(10)}$$

where $\tilde{y}_{it}, \tilde{z}_{it}$, and $\tilde{e}_{it}$ are typical elements of, respectively, $\tilde{y}_i = M_i y_i$, $\tilde{Z}_i = M_i Z_i$, and $\tilde{e}_i = M_i e_i$; $M_i = (I_i - X_i(X_i'X_i)^{-1}X_i')$; and I_i denotes the $(T_i \times T_i)$ identity matrix. Although (10) is an equivalent representation of the model in (9),

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the former is more intuitive, when it is generalized toward a nonlinear relationship depending on predetermined macroeconomic state variables, i.e.,

$$\tilde{y}_{it} = \tilde{z}_{it}'\gamma_i(\omega) + \tilde{e}_{it},\ \omega_{it} = R_{it-1}, \pi_{it-1}, (R_{it-1} - \pi_{it-1}), \Delta R_{it-1}, \Delta\pi_{it-1}, \Delta(R_{it-1} - \pi_{it-1}). \quad \textbf{(11)}$$

As it allows for nonlinear dynamics, the specification in (11) is similar to a recent proposal by Escribano (2004) formalizing the EC parameter in terms of a spline function. As a particular distinction, the variant in (11) requires that potential forces behind the EC dynamics can be attributed to measurable economic states ω. Choosing some function of time, $\omega_{it} = t/T$, say, might be seen to closely approximate the spline smoothing approach within the functional coefficient framework. With $K(\cdot)$ and h denoting a kernel function and the bandwidth parameter, respectively, and $K_h(u) = K(u/h)/h$, the semiparametric Nadaraya–Watson type estimator [Nadaraya (1964); Watson (1964)] is

$$\hat{\gamma}_i(\omega) = \mathcal{Z}_i^{-1}(\omega)\mathcal{Y}_i(\omega), \quad \textbf{(12)}$$

where $\mathcal{Z}_i(\omega) = \sum_{t=1}^T \tilde{z}_{it}\tilde{z}_{it}' K_h(\omega_{it} - \omega),\ \mathcal{Y}_i(\omega) = \sum_{t=1}^T \tilde{z}_{it}\tilde{y}_{it} K_h(\omega_{it} - \omega)$.

3.2. Functional Models

As is typical in nonparametric regression, local estimates as (12) suffer from a trade-off between bias and variance. Owing to the use of a kernel function, nonparametric estimates may be seen as local averages of the underlying function, such that $\hat{\gamma}_i(w)$ is essentially an estimate of a smoothed version of $\gamma_i(w)$. Because we are interested in the overall cross-sectional behavior of functional coefficients, our main conclusions are likely unaffected by (small) local biases.

Asymptotic results in Cai et al. (2000) are derived for (auto)regression designs with stationary variables. As is known from parametric cointegration modeling, estimates of the EC parameter behave as coefficient estimates in stationary models whenever the involved levels of interest rates and inflation are both stationary or both nonstationary but cointegrated. Furthermore, taking results in Karlsen et al. (2007) into account, we consider the model in (11) as a local approximation of the ECMs in (4) and (5) for those cross-section members where common degrees of integration are diagnosed for inflation and nominal interest rates ["$I(0)$" or "$I(1)$"]. With respect to the DOLS regression, stationarity is clearly an unsuitable assumption for the major fraction of the investigated cross section(s). However, from recent work on the statistical properties of nonparametric regressions with integrated processes [Park and Phillips (2001); Karlsen et al. (2007)], the functional coefficient estimates $\hat{\theta}_i^{(\bullet)}(\omega)$ are known to share typical properties of nonparametric estimators. In particular, owing to results in Park (2005), $\hat{\theta}_i^{(\bullet)}(\omega)$ converges to $\theta_i^{(\bullet)}(\omega)$ and shows some (local) bias in case of a finite bandwidth parameter.

3.3. Bandwidth Selection

The choice of the bandwidth parameter is of crucial importance for the factor-dependent estimates in (12). Choosing too small a bandwidth may result in a wriggly pattern of the semiparametric estimates. If h is prohibitively large, $\hat{\gamma}_i(\omega)$ degenerates to time-invariant EC or Fisher coefficient estimates. In this sense, the semiparametric approach nests the common cointegration analysis, and it is interesting to see if local averaging supports the parametric model. We choose the bandwidth parameter locally as

$$h_i(\omega) = a\hat{f}_i(\omega)^{-0.25}, \tag{13}$$

where $\hat{f}_i(\omega)$ is a kernel density estimate for the factor sample ω_{it}, $t = 1, \dots, T_i$. With respect to the parameter a we distinguish two alternative settings, $a = 2T_i^{-0.25}$ and $a = T_i^{-0.25}$, for which the factor corresponds to level variables and first differences, respectively. The former choice turned out to be preferable, as the smaller bandwidth selection with $a = T_i^{-0.25}$ delivered somewhat wriggly functional estimates when conditioning on level factors. Jennen-Steinmetz and Gasser (1988) show that for stationary factor variables the upper choice of the local bandwidth roughly corresponds to spline smoothing. To implement density estimation for ω_{it} we use $h = 1.06\sigma_\omega T_i^{-0.2}$ as the bandwidth parameter [Silverman (1986)], where σ_ω is the unconditional standard deviation of the factor. Throughout, $K(u)$ is the quartic kernel, $K(u) = 15/16(1-u^2)^2 I(|u| < 1)$.

3.4. Conditioning Factors

As potential economic states for conditional modeling, we consider the lagged levels of nominal interest ($\omega_{it} = R_{it-1}$), inflation ($\omega_{it} = \pi_{it-1}$), or real interest rates ($\omega_{it} = R_{it-1} - \pi_{it-1}$) and the corresponding first differences [$\omega_{it} = \Delta R_{it-1}$, $\Delta\pi_{it-1}$, $\Delta(R_{it-1} - \pi_{it-1})$]. Note that in absolute values the latter factors could be interpreted as an approximation of (real) interest rate or inflation risk. Amato (2005) summarizes the relevance of the real interest rate in many theories explaining business cycles or inflation fluctuations. Moreover, the real interest rate is often a core building block for determining the natural rate of interest, which is the real short-term interest rate that is consistent with output growing at the natural rate and constant inflation. The stance of monetary policy is effectively expansionary (contractionary), if the shor-term real interest rate is below (above) the natural rate. Adopting a Kalman filter approach, Laubach and Williams (2003) estimate the latent natural rate to vary between 2% and 4% in the United States. Recent results for the euro area point in the same direction [ECB (2004)]. Accordingly, real rate values below 1% (in excess of 5%) could be seen as indicating an expansionary (contractionary) stance of monetary policy.

Because we are interested in the cross-sectional characteristics of local estimates $\hat{\gamma}_i(\omega)$, the functional model is evaluated at a grid

$$\omega_k = \omega_{min} + g_\omega k, \quad k = 0, 1, 2, \ldots, 51, \tag{14}$$

where the factual support is factor-specific and depends on the choice of ω_{min} and g_ω. Equidistant grid points are determined over the following ranges of level variables: $0\% \le R_{it-1} \le 20\%$ (nominal interest rate), $0\% \le \pi_{it-1} \le 20\%$ (inflation rate), $-10\% \le R_{it-1} - \pi_{it-1} \le 10\%$ (real interest rate). If rate changes are used for conditioning, the relevant support is $-2\% \le \Delta\omega_{it} \le 2\%$, $\omega_{it} = R_{it-1}, \pi_{it-1}, (R_{it-1} - \pi_{it-1})$.

Available time series enter the semiparametric estimator in weighted form. Functional estimates are "most representative" for factor observations in the neighborhood of local points ω_k. For this reason it is important to characterize the empirical distribution of the employed factor variables. Figure 1 illustrates some features of the empirical distributions of factor variables for the entire sample period and the considered subperiods. The single graphs in Figure 1 characterize the empirical distribution of factor variables over ten equidistant subintervals. Thus, in case level (change) variables are used as factors, each bar in Figure 1 corresponds to factor observations over an interval of 2 (0.4) percentage points. Most results in Figure 1 are provided for the full sample period (1960–2004) and the cross section characterized by jointly nonstationary inflation and interest rates ["$I(1)$"].

The first row in Figure 1 depicts the fraction of the 57 "$I(1)$" economies for which observations within the specified support are available. These fractions mostly exceed 90%. Thus, the selected factor support is "relevant" for a rather substantial fraction of the cross section. With regard to the level of short-term interest rates (real rates), it holds that very small values between 0% and 2% (-10% to -8%) are less typical, and are quoted for $\approx$14% ($\approx$58%) of the 57 economies. The second "line" in Figure 1 displays the fractions of panel observations falling within the specified intervals of the factor support. The empirical distributions of level factors have their modes between 4.0% and 6.0% (interest rates), 2.0% and 4.0% (inflation rates), and 0% and 4.0% (real rates). Notably, even the tails of the empirical factor support comprise considerable fractions of the available factor quotes. For instance, from almost 19,000 observations [1960–2004, "$I(1)$"], interest rates between 0% and 2% and 18% and 20% have empirical frequencies of 0.021 and 0.029, respectively. The empirical distributions of differenced factor variables are displayed in the third row of Figure 1. These distributions are "more symmetric" in comparison with the level factors and reflect factor stationarity. The most likely adjustments of these rates are between -0.4 and +0.4. Adjustments of interest rates within this range have a frequency of about 83%. For comparison, the distribution of real rate or inflation adjustments is more spread, because absolutely larger changes show higher frequencies, as displayed for interest rate changes. The bottom line in Figure 1 illustrates that higher real rates are more frequent during

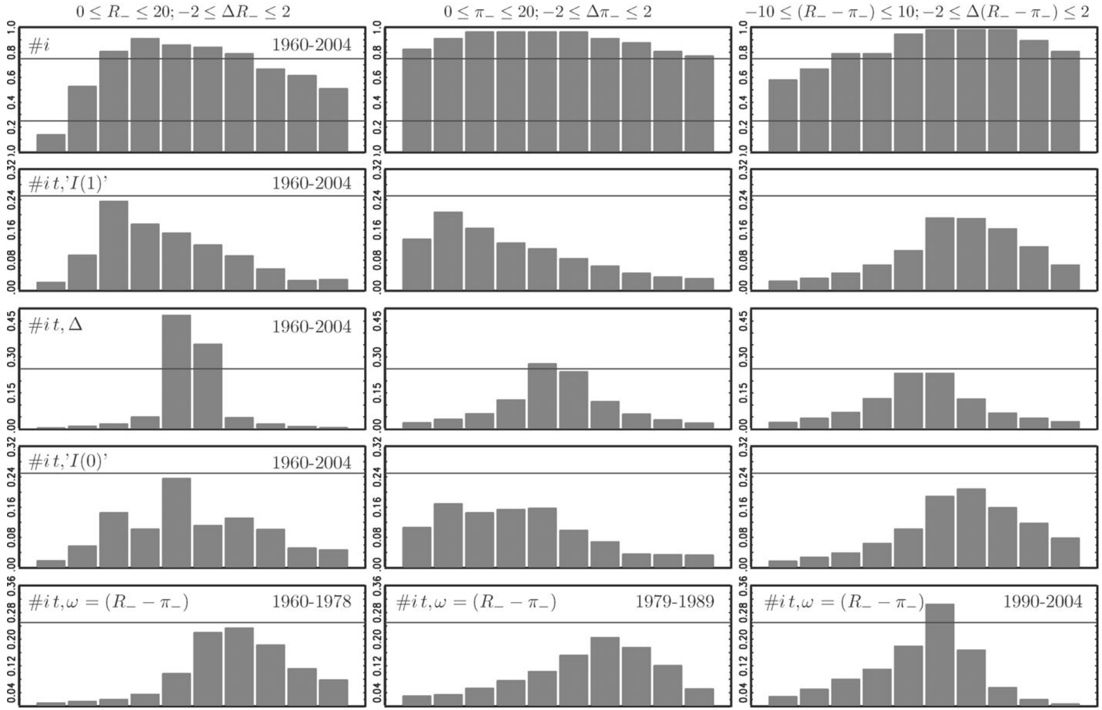

FIGURE 1. Distributional characteristics of state variables: Fractions of cross-section members ["$I(1)$," 1960–2004], for which indicated observations are available (first line), fractions of available observations (level factors second line, first differences third line, real rate by subsamples fifth line, "$I(0)$" cross section, 1960–2004, fourth line). The x-axis represents a subdivision of the factor-specific support in ten equidistant bin; each bin covers 2% and 0.4% for factors in levels and first differences, respectively. To facilitate comparability of bar plots, the horizontal lines indicate 25% levels in all panels and, additionally, the 75% level in the top row.

the first subperiod (1960–1978), whereas in contrast the time span from 1990 to 2004 is characterized by relatively low real rates. During the period 1979–1989 the empirical distribution of real rates features the highest dispersion over the cross section. This period marks the beginning of disinflation policies in industrial economies. Policy changes were established at diverse time periods and with distinct emphasize. Hence, the real interest rates diverge to a larger extent in comparison with the other two subperiods [Desroches and Francis (2007)]. Finally, the fourth line in Figure 1 displays the empirical support of level factors for the overall sample period and conditional on the "$I(0)$" cross section. This cross section comprises only 13 economies, so that direct comparability with the second line of Figure 1 is weakened. The provided histograms for real rate levels are, however, remarkably invariant over both cross sections, "$I(0)$" and "$I(1)$." The distribution of interest rates is more dispersed for the "$I(0)$" sample and its mode levels are shifted to the right (between 8.0% and 10%).

3.5. Inference for Functional Coefficient Models

To obtain pointwise confidence intervals for $\hat{\gamma}_i(\omega_k)$, resampling techniques as outlined in Cai et al. (2000) are widely applied. For a given member of the cross section, pointwise confidence bands offer some guidance to evaluate the adequacy of linear models as specified in (4) to (6). Instead we base inferential issues on the variation of semiparametric estimates over the cross section. The panel structure of the sample is used to characterize the mean group level of the parameters of interest. Throughout, functional estimates are provided, graphically showing that $\bar{\gamma}(\omega_k) \pm 2\sigma(\bar{\gamma}(\omega_k))$, where $\sigma(\bar{\gamma}(\omega_k))$ is the estimated standard deviation of $\bar{\gamma}(\omega_k)$. Note that under cross-sectional parameter homogeneity and consistent estimation, such a confidence band covers the true parameter with a probability of approximately 95%.

4. FUNCTIONAL ESTIMATES

In this section we discuss semiparametric functional estimates. First, we add a few details of model implementation and the selection of empirical results that are discussed in more detail. Then significance or time variation of both EC dynamics and the Fisher coefficient are addressed.

4.1. Preliminary Remarks

Because functional estimates are most suitably provided graphically the full set of results [two cross sections—"$I(0)$" and "$I(1)$"—four sample periods and six factor variables] cannot be displayed because of space considerations, but is available from the authors upon request. It turned out that state-dependent patterns of EC parameters or the Fisher coefficient are less pronounced for the cross section of stationary ["$I(0)$"] in comparison with nonstationary systems ["$I(1)$"]. Moreover, the latter cross section includes considerably more economies for all subperiods than the former. Functional results for the overall sample period are similar to some average pattern obtained by (nonlinear) weighting of subsample-specific estimates. Therefore, functional results are only provided for the "$I(1)$" cross section and the three subsample periods.

The following discussion of semiparametric estimates first addresses state dependence of cointegration properties by evaluating mean group functional EC parameters $\bar{\alpha}_1(\omega_k)$. Then weak exogeneity of inflation is briefly discussed by considering local estimates $\bar{\alpha}_2(\omega_k)$. Finally, conditional on states that are most likely characterized by existing equilibrium relationships, functional estimates of the Fisher coefficient $[\bar{\theta}(\omega_k)]$ are discussed.

4.2. Error Correction Dynamics

Figure 2 displays factor-dependent estimates of the EC coefficient governing interest rate adjustments. The conditioning factors are the lagged nominal interest ($\omega = R_-$) and inflation rate ($\omega = \pi_-$) and first differences ($\omega = \Delta R_-,\ \Delta\pi_-$). From the left- to the right-hand side the columns of Figure 2 show estimates for the periods 1960–1978, 1979–1989, and 1990–2004, respectively.

The functional mean group EC coefficient estimates are mostly significantly negative when conditioned upon $\omega = \pi_-$ or $\omega = R_-$. The overall evidence in favor of an equilibrium relationship linking inflation and interest rates is particularly weakened for states of high (lagged) nominal interest or inflation rates ($R_-,\ \pi_- > 15\%$). Such states are likely characterized by failure or uncertainty of financial investments to earn sufficiently high real returns. Noteworthily, the unconditional average Fisher coefficient is $\approx$0.3 and $\approx$0.7 (see Table 1) for the subperiods 1960–1989 and 1990–2004, respectively. In states of high nominal interest or inflation rates, an established linkage between nominal interest and inflation rates is distorted and stronger (i.e., even positive) adjustments of interest rates might occur to satisfy the public's claim for real returns. In a similar fashion evidence against cointegration is documented for states of relatively large upward changes of inflation ($\Delta\pi_- > 1.2\%$). Because such states most likely coincide with scenarios of high inflation rates, this result reflects the former argument that the public is not willing to accept real losses in states of accelerating inflation. Moreover, Ragan (1994) documents a positive relation between the inflation rate and its standard deviation for 22 OECD countries. Similar evidence is confirmed for the G7 in Berument et al. (2007). In scenarios of high level and risk of inflation, increasing and potentially nonstationary risk premia could weaken the equilibrium relation between inflation and interest rates.

Periods of higher interest rate changes go along with a strengthening of EC dynamics. Note that large positive adjustments of policy rates are likely an indication of more restrictive monetary policies. As the latter diagnosis is most evident for the subperiods 1979–1989 and 1990–2004, it appears that the credibility of monetary authorities has been strengthened over these decades.

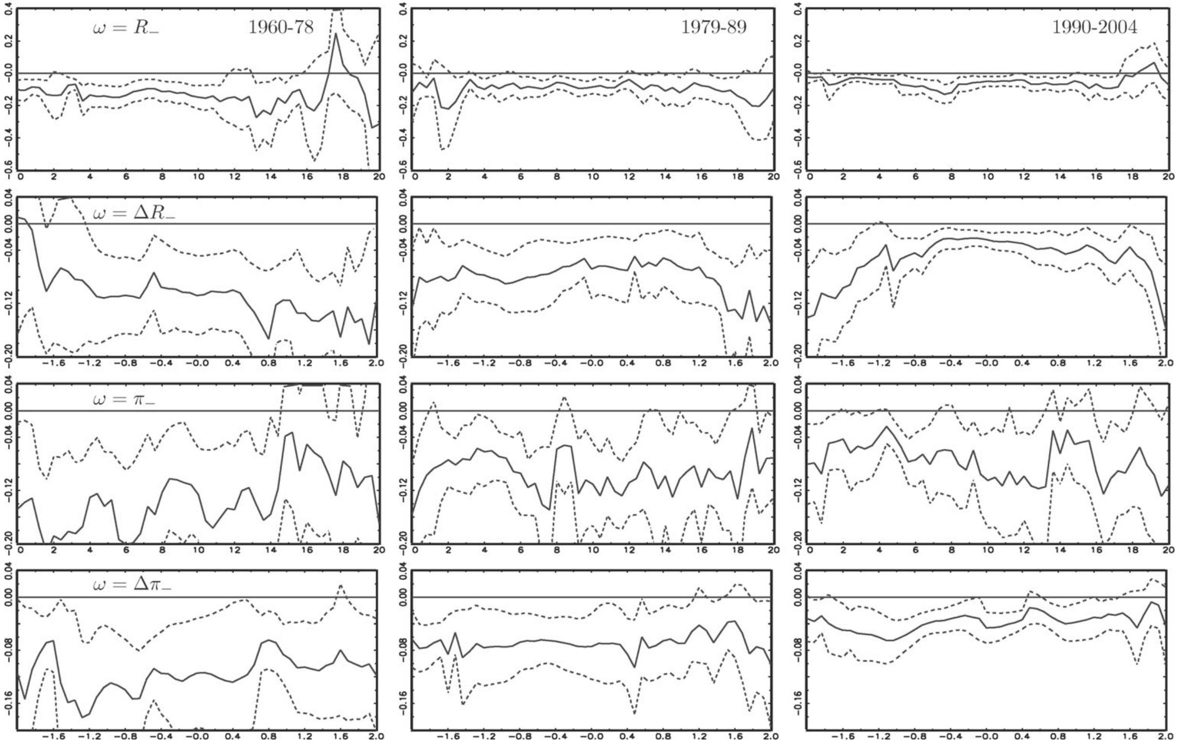

FIGURE 2. Functional mean group estimates of EC coefficient governing the interest rate adjustments over three subperiods. Factor variables are lagged level and first differences of interest rates (lines 1 and 2) and the lagged level and first differences of inflation. Dashed lines correspond to an approximate 95% confidence interval. For better reference, a zero-level horizontal line is also shown. "Uninformative" lower bounds of confidence intervals might be cut to improve the scale of the graphs, which is identical within particular lines.

EC dynamics conditioned upon real rate factors [$\omega = (R_{-} - \pi_{-}),\ \Delta(R_{-} - \pi_{-})$] are displayed in the first two rows of Figure 3. With regard to the real rate level, the EC parameter is mostly significantly negative for $(R_{-} - \pi_{-}) > -2\%$. Thus, for a stable link between nominal interest and inflation rates, the real rate has to be sufficiently nonnegative. To underpin this argument, it is noteworthy that periods of low real rates often coincide with states of high inflation that have been discussed before. In scenarios where the real rate is below -2%, monetary policy is presumably expansionary [Amato (2005)]. The public suffers from real losses that it is not willing to accept over the medium or longer term. Functional estimates conditioning on real rate changes confirm this perspective. Negative adjustments in the real rate entail the risk that the public might call for excess compensation in comparison with the historically experienced linkage of nominal interest and inflation. Evidence in favor of cointegration is the strongest over states of positive real rate adjustments. In these situations the monetary policy of a central bank becomes more tightening, setting, for instance, the short-term interest rate in excess of the actual inflation rate.

4.3. Weak Exogeneity

Assessing local characteristics of weak exogeneity, we find that adjustments of the inflation rate in response to violations of the (potential) equilibrium relation are almost uniformly insignificant. Therefore, we do not provide functional estimates $\bar{\alpha}_2(\omega_k)$ in detail. The only states identified as showing significant stabilizing responses of inflation are obtained by conditioning on interest rates or first differences thereof. The lower rows of Figure 3 show functional mean group estimates $\bar{\alpha}_2(\omega)$ for $\omega = R_{-}$ and $\omega = \Delta R_{-}$. For the period of inflation targeting (1990–2004), it turns out that under small interest rate adjustments (or "stable" interest rates), functional EC parameters are significantly positive. This result underpins the improved credibility of monetary authorities, because in these situations the transmission of monetary policy appears to operate additionally through the inflation expectations channel. The public believes that monetary authorities have superior information about the future development of inflation and act in line with the future inflation rate. In such states central banks are able to reduce the interest rate and likely signal the future inflation rate to be lower than the actual one [Miskhin (2003, Chapter 28)]. This mechanism, however, necessitates a high credibility of the monetary authorities.

4.4. The Fisher Coefficient

Figure 4 illustrates mean group Fisher coefficients conditioned on $\omega = R_{-},\ \Delta R_{-}, \pi_{-},\ (R_{-} - \pi_{-})$. Conditioning on first differences $\omega = \Delta\pi_{-},\ \Delta(R_{-} - \pi_{-})$ obtained almost constant coefficient patterns and, therefore, graphical results are not displayed, to economize on space. The unconditional levels of the estimated Fisher coefficient roughly correspond to the parametric mean group

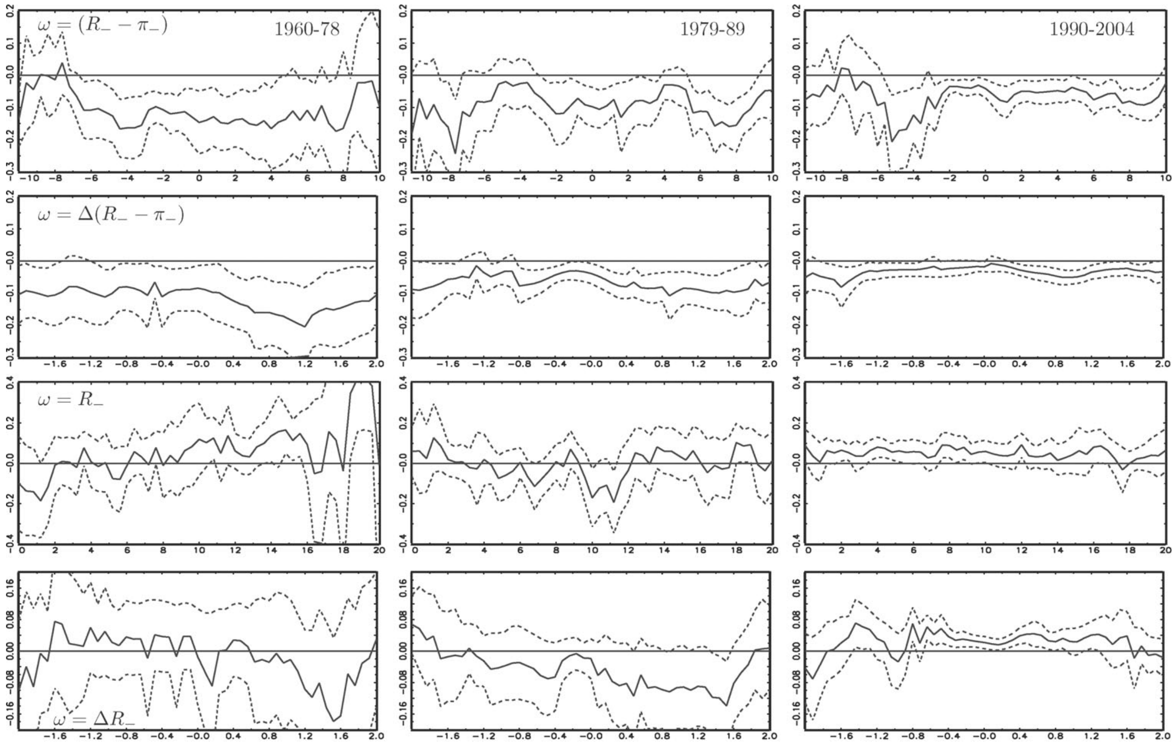

FIGURE 3. Functional mean group estimates of EC coefficient governing the interest rate adjustments conditional on lagged real interest rate levels and first differences (upper two lines). The lower lines display functional patterns of the inflation rate adjustment parameter $[\alpha_2(\omega_k)]$ conditional on lagged interest rate levels and first differences thereof. For further notes see Figure 2.

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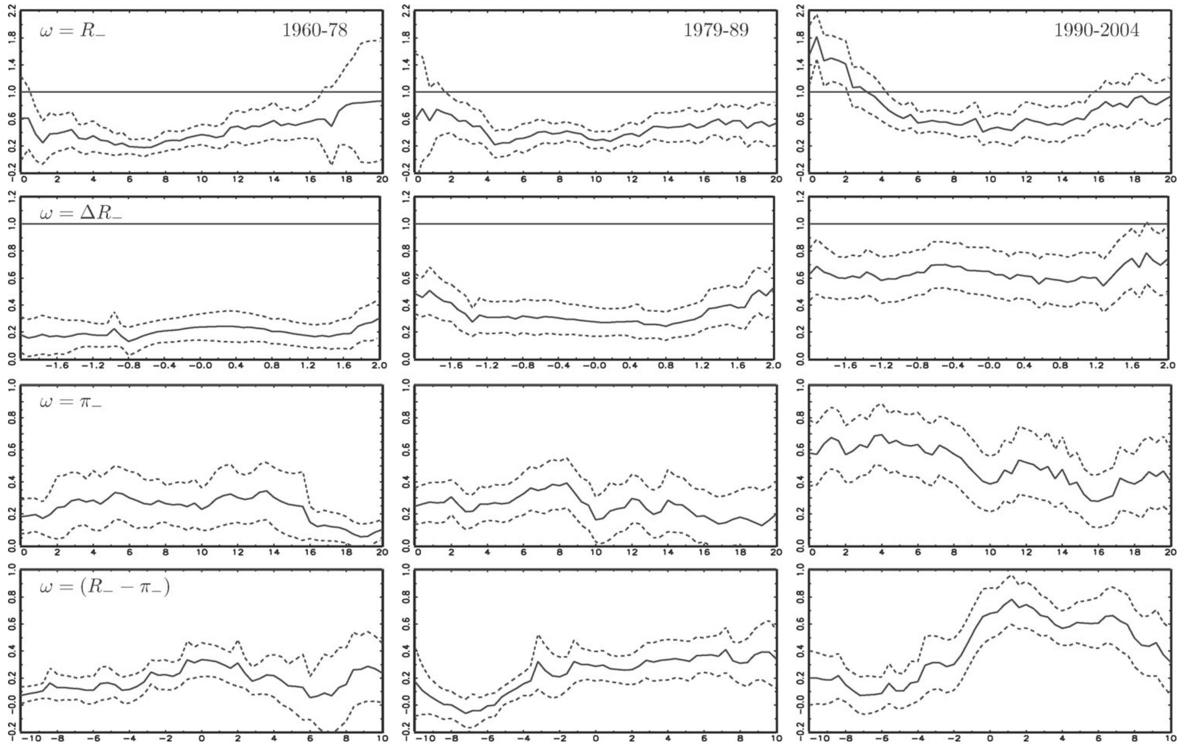

FIGURE 4. Functional mean group estimates of the Fisher coefficient conditioning on three alternative (lagged) level factors and first differences of interest rate (second line). Local averaging applied only for estimates $\hat{\theta}_k \in [-1, 3]$ to guard against adverse impacts of outlying estimates. The prevalence of local outliers is similar to the counts given for the parametric DOLS estimator in Table 1. Eventual horizontal lines indicate a level of unity. For further notes see Figure 2.

estimates provided in Table 1. For the discussion of the mean group long-run parameter, it is natural to concentrate on economic states that are characterized by a stable equilibrium relationship. According to the results for the functional EC parameters, cointegration is most likely over states with medium interest rates, small or moderate inflation, and positive real interest rates.

Over states with medium interest rates or low to medium inflation rates, the mean group Fisher coefficient is almost stable. Throughout, it is significantly positive and smaller than unity, which is at odds with a stabilizing policy under the notion of a Taylor-type reaction function of monetary authorities. As already documented in Table 1, the Fisher coefficient is the largest for the period 1990–2004 [$\bar{\theta}(\omega_k) \approx 0.6$], whereas corresponding estimates for the two former periods are markedly smaller [$\bar{\theta}(\omega_k) \approx 0.25$: see also Table 1].

With regard to the strategy of monetary policy, we diagnose that the Fisher coefficient increases over states of large positive interest rate adjustments. Thus, the unconditional mean-group estimates given in Table 1 are not uniformly representative for all economic states. However, even in these states, the parameter remains significantly below unity. Interestingly, for the inflation-targeting regime and conditioning on high interest rates ($R_- > 16\%$) or interest rate adjustments ($\Delta R_- > 1.6\%$), the Fisher coefficient differs only insignificantly from unity. Interpreting the Taylor rule as a reaction function of the central bank, these states are characterized by active central bank interest rate reactions to the development of the inflation rate to stabilize the price level. In this sense, the strongest policy reactions are diagnosed for the last subperiod and states of low interest rates that obtain local mean group Fisher coefficients in excess of unity. Both findings strengthen the importance of the real rate for the implementation of monetary policy and conform with Rogoff (2003) pointing out that central banks pay more attention to price stability during the last decades.

Regarding the lower right panel of Figure 4, we observe a hump-shaped functional Fisher coefficient conditional on positive real rates that characterize the inflation-targeting regime. Again, the Fisher coefficient is significantly less than unity. In periods of strongly expansionary or contractionary monetary policy, the linkage of interest and inflation rates becomes unstable. Interestingly, however, it is closest to unity for real rates around 1%. The impression that the Fisher coefficient diminishes over states of high real rates might reflect situations of successful anti-inflationary policies. If a central bank has gained sufficient credibility, the short-term interest rate could diminish while leaving the inflation rate almost unchanged.

The largest functional Fisher coefficient estimates are found for states of low positive real rates and, similarly, low nominal interest rates. In particular, for the last subsample, a strongly expansionary monetary policy goes along with marked deviations from the Fisher relation. This instability of the long run parameter confirms the evidence in Amato (2005). He documents for the United States and the United Kingdom, the sensitivity of the policy rate coefficients in a Taylor rule–type equation to variations in the long-run natural rate.

5. CONCLUSIONS

In this study, we investigate the link between inflation and nominal interest rates for a cross section of 114 economies over a period covering at most 45 years of monthly observations. For the full sample period, it is diagnosed that for 44 economies nominal interest and inflation rates are characterized by distinct degrees of integration, which is at odds with the Fisher hypothesis. We focus on a general country and time-specific semiparametric approach, supporting the view that in nonstationary systems, cointegration features are likely time-varying.

Error correction dynamics and the Fisher coefficient are found to be heterogeneous over the time and the cross-section dimension. The empirical Fisher coefficient is mostly less than unity. Hence, the analysis makes it possible to characterize economic conditions where a basic economic relationship such as the often (implicitly) presumed stability of real interest rates fails. Interest rates and inflation are found to exhibit a long-run equilibrium relation for numerous economic states. Moreover, in states of negative real interest rates, a long-run equilibrium relation may not exist. Interestingly, for the period of inflation targeting, the violations of the long-run relationship invoke stabilizing adjustment of the inflation rate in scenarios of (almost) stable interest rates. From this, one may conclude that there has been a general improvement of the credibility of monetary authorities.

In summary, the worldwide evidence in favor of a stationary real short-term interest rate is weak. The long-run reaction of central banks to inflation is weaker than necessary to ensure a stable and positive real interest rate. Moreover, the average Fisher coefficient is not time-invariant. Interestingly, the world's average Fisher coefficient is smaller over earlier sample periods. One explanation might be that the policy objective of some central banks was not only to maintain price stability but also to foster output growth, as for example in the United States. Similarly, the targeting of exchange rates, as typical for the Bretton Woods era, could explain empirical Fisher coefficients below unity. In addition, globalization might keep inflation low so that the weak reactions of central banks may not carry an immediate risk of accelerating price levels. Even for the most recent subperiod, characterized by widespread inflation-targeting policies, the mean group Fisher coefficient is (mostly) smaller than unity. A few selected states are identified for which the Fisher hypothesis is not rejected and the Taylor rule might obey the interpretation of a policy reaction function. Given the risk that the short-term real interest rate as a core variable in New Keynesian monetary models fails stability, inflation fighting might deserve (re)considering the role of monetary aggregates and exchange rate and wage developments or of asset prices. We regard these most likely state-dependent impacts as an issue of future research.

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APPENDIX: LIST OF COUNTRIES

1. *Members of the first subsample, 1960–1978*: Austria, Bahamas, Barbados, Belgium, Botswana, Burkina Faso, Burundi, Cameroon, Canada, Colombia, Costa Rica, Cote d'Ivoire, Denmark, Egypt, Finland, France, Gabon, Germany, India, Italy, Jamaica, Japan, Jordan, Kenya, Korea, Malaysia, Mexico, Netherlands, Niger, Nigeria, Pakistan, Peru, Philippines, Senegal, Singapore, South Africa, Sweden, Switzerland, Thailand, Togo, Trinidad, Tunisia, United Kingdom, United States of America.

2. *Additional cross-section members entering the second subsample, 1979–1989*: Argentina, Bahrain, Brazil, Central African Republic, Chad, Democratic Republic of Congo, Cyprus, Ecuador, Fiji, Gambia, Ghana, Greece, Guinea Bissau, Honduras, Hungary, Iceland, Indonesia, Israel, Kuwait, Mali, Malta, Mauritius, Nepal, Netherlands Antilles, Nicaragua, Norway, Portugal, Seychelles, Sierra Leone, Solomon Islands, Spain, Sri Lanka, Swaziland, Tanzania, Turkey, Uganda, Uruguay, Zambia, Zimbabwe.
3. *Additional cross-section members entering in last subsample, 1990–2004*: Algeria, Armenia, Bangladesh, Benin, Bolivia, Bulgaria, Chile, China Hong Kong, Republic of Congo, Croatia, Czech Republic, Dominica, Dominican Republic, El Salvador, Estonia, Georgia, Grenada, Guatemala, Guyana, Haiti, Kazakhstan, Laos People's Democratic Republic, Luxembourg, Macedonia, Mongolia, Myanmar, Namibia, Paraguay, Poland, Romania, Slovak Republic, Slovenia.