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What the Yield Curves say About Inflation: Does it Change Over Time?

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ABSTRACT/RESUME

The paper investigates the information content of yield curves regarding future inflation, using the example of the G-7 countries. The empirical results show substantial variation of results across countries, and a significant information content is identified for the United States, the United Kingdom, Germany and Canada. The results also vary with the choice of the sample period. Nevertheless, the relationship appears to be structurally stable in Germany and Canada. By contrast, there is evidence for structural instability in the United States and the United Kingdom, possibly reflecting changes in their monetary policy regimes.

Keywords: Monetary policy indicator, information content of term structure of interest rates, structural stability test, monetary policy regimes.

JEL Classification System-Numbers: E42, E43.

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En prenant comme exemple les 7 grands, ce document examine le contenu informatif des courbes des rendements en regard à l'inflation à venir. Les résultats empiriques montrent des résultats dont les variations sont substantielles d'un pays à l'autre. On peut identifier un contenu informatif significatif dans le cas des États-Unis, du Royaume-Uni, de l'Allemagne et du Canada. Ces résultats varient aussi en fonction du choix de la période de référence. Néanmoins les relations semblent être structurellement stables en Allemagne et au Canada. Par contre, on trouve aux États-Unis et au Royaume-Uni des preuves d'instabilité structurelle qui peuvent refléter des changements dans leur régime de politique monétaire.

Mots-clé : Indicateurs de politique monétaire, contenu informatif en terme de structure des taux d'intérêt, test de stabilité structurelle, régimes de politique monétaire.

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WHAT THE YIELD CURVES SAY ABOUT INFLATION: DOES IT CHANGE OVER TIME?

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1. Introduction*

Yield curves or term structures of interest rates are one of the financial market indicators used by monetary policy makers to extract information on future interest rate and inflation developments. While the extent and sophistication of the uses of yield curves to extract information vary across central banks and over time,[1] the forward looking nature of underlying bond prices has meant that monetary policy makers monitor the slope of yield curves as a general rule. Often, the information content of the term structure about inflation is defined as the ability of the slope of the term structure to predict changes in inflation rates. Specifically, if real interest rates and term premia are constant over time, the difference between short-term and long-term interest rates should be a linear function of expected inflation changes.

The relationship under investigation is likely to be subject to variation as a result of changes in the structure of the economy. In this context some results are available. For example Mankiw and Miron (1986) find that the term structure in the United States was much more useful in predicting future interest rate changes before the founding of the Federal Reserve System. Mishkin (1990*a*, *b*) finds that the U.S. term structure is more informative about inflation in the period before the Fed's announcement of a change

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1. A recent account of the use in practice of financial market indicators by central banks is provided in Mylonas and Schich (1999).

in its monetary policy operating procedures, in October 1979. Another example is the choice of the exchange rate regime; Gerlach and Smets (1997) suggest that exchange rate pressures in regimes with pegged exchange rates, have obscured the information in yield spreads about future interest rate changes. Recently, attention has been drawn to the effects of the regulation of financial markets and the institutional structures of government bond markets on the information contained in asset prices. For example, Bernard and Gerlach (1996) find that the information content of yield spreads regarding future activity is higher in Canada, Germany, and the United States than in Japan, and argue that this reflects the tighter regulation of financial markets in the latter. De Broeck, Guillaume and Van der Stichele (1998) find that the introduction of futures and options on government bonds tended to improve the informational efficiency of the prices of the underlying bonds.

These changes in the structure of the economy, i.e. changes in the monetary policy or exchange rate regime, financial market regulation or the institutional structure of bond markets, may also imply variation and instability in the relationship between yield spreads and future inflation. Several results from previous empirical studies are available in this context, showing that the information content of the term structure is subject to change over time. For example, Jondeau and Ricart (1997) fail to find any significant information content in the French term structure using their full sample, but do so on the basis of a sub-sample (Table 1). Koedijk and Kool (1995) also consider two different sample sizes for Germany and find evidence for a significant information content only for the larger sample, though the evidence is very weak. By contrast, Gerlach (1997) and Schich (1996) find strong evidence for a significant information content in the German term structure, each using a considerably larger sample size than Koedijk and Kool. While Gerlach points out that *small* variations of the sample size do not “materially affect” his results, the combined evidence suggests that this does not necessarily hold for *large* sample variations. Against the background of all these results, a study of the stability of the predictive content of the term structure about inflation appears useful.

The present paper follows the methodology of the papers in the tradition of Mishkin, defining the information content of the yield curve as the ability of its slope to predict future changes in inflation. It differs from them in two respects. First it focuses on the issues of time variation and stability of results. The paper addresses the question of whether the relationships are stable over time, or exhibit a structural break. Specifically, the techniques of Andrews (1993) and Ghysels, Guay, and Hall (1997) are applied to test for structural change in models estimated via the Generalised Method of Moments (GMM). These techniques allow the instability to occur at an unknown point. This is appropriate given that the exact timing of the effect of the above mentioned factor changes on the structure of the economy is generally uncertain. The results are also compared with those obtained from direct tests for breakpoints, where the structural breaks are assumed to be known *a priori*. Second, the present paper uses a much longer and more recent sample period than previous studies.[2]

2. Methodology

2.1 The definition of information content

The methodological approach to testing for the information content follows Mishkin (1990*a*, *b*, 1991), who defines the information in the term structure as the ability of the slope to predict changes in the inflation rate. The starting point is a Fisher decomposition of the nominal zero-coupon yield:

$$z_t^j = rr_t^j + E_t\left[\pi_t^j\right], \tag{1}$$

2. E.g. the data samples are longer and more recent than those used in Browne and Manasse (1990), Day and Lange (1997), Gerlach (1997), Jondeau and Ricart (1997) Jorion and Mishkin (1991), and Koedijk and Kool (1995), and Mishkin (1990a, b, 1991) and Schich (1999). This is important as it increases the power of the stability tests, and allows us to explicitly address the issue of sample dependency of results using rolling regressions. Nevertheless, the sample for France, Italy and Japan are relatively short. The data consist of end-of-month interest rates from one to five years for Canada (starting January 1967), France (January 1980), Germany (January 1967), Italy (October 1985), Japan (March 1980), the United Kingdom (January 1970) and the United States (June 1961), all series ending in 1998. They are obtained from national central banks and the beginning dates of the series are dictated by data availability.

where z_t^j is the j-year zero-coupon yield, E_t is the expectations operator conditional on information available at time *t*, rr_t^j is the *j*-year *ex ante* real interest rate and π_t^j is the actual forward inflation rate over the next *j* years, which is computed as $\left(\left(P_{t+12j} / P_t\right)^{1/j} - 1\right) * 100$ with P_t denoting the consumer price index in month *t*. Assuming rational expectations the *realised* inflation rate over the next *j* years can be written as the *expected* inflation rate plus a serially uncorrelated, zero-mean error ε_t^j ,:

$$\pi_t^j = E_t\left[\pi_t^j\right] + \varepsilon_t^j \ , \tag{2}$$

where $\varepsilon_t^j = \pi_t^j - E_t\left[\pi_t^j\right]$ is the expectation error of inflation. Substituting for $E_t\left[\pi_t^j\right]$ from equation (1) we obtain

$$\pi_t^j = z_t^j - rr_t^j + \varepsilon_t^j \ . \tag{3}$$

Hence the realised inflation rate over the next *k* (with *k*< *j*) years can be expressed as:

$$\pi_t^k = z_t^k - rr_t^k + \varepsilon_t^k \ . \tag{4}$$

Subtracting (4) from (3), we obtain an expression of the changes in the realised inflation rate between the two periods *j* and *k*:

$$\pi_t^j \ - \ \pi_t^k = -\left(rr_t^j - rr_t^k\right) + \left(z_t^j - z_t^k\right) + \left(\varepsilon_t^j - \varepsilon_t^k\right). \tag{5}$$

Assuming that the differences between the real interest rates for maturities *j*, *k* equal some constant $\alpha^{j,k}$ plus a zero mean random variable $\upsilon_t^{j,k}$, specifically $-\left(rr_t^j - rr_t^k\right) = \alpha^{j,k} + \upsilon_t^{j,k}$, equation (5) can be written in estimable form:

$$\pi_t^j \ - \ \pi_t^k = \alpha^{j,k} + \beta^{j,k}\left(z_t^j - z_t^k\right) + u_t^{j,k} , \tag{6}$$

where $u_t^{j,k} = \upsilon_t^{j,k} + \left(\varepsilon_t^j - \varepsilon_t^k\right)$. Given rational expectations, and that no shift in the inflation regime at some future date is expected (Evans and Wachtel, 1993), the expectation errors ε_t^j and ε_t^k are orthogonal to the right-hand regressors. In other words, expectation errors are indeed unexpected, conditional on all available information at time t including z_t^j and z_t^k -- a necessary condition for consistent estimation (Browne and Manasse, 1990). If the slope of the real term structure $\left(rr_t^j - rr_t^k\right)$ is constant, then $\upsilon_t^{j,k}$ is equal to zero and the nominal term structure $\left(z_t^j - z_t^k\right)$ is the optimal predictor of $\left(\pi_t^j - \pi_t^k\right)$. If the slope of the real term structure is not constant because $\upsilon_t^{j,k}$ is predictably different from zero, then the GMM estimate of $\beta^{j,k}$ still has a probability limit of one, however, the nominal term structure may not be the optimal predictor of future inflation changes.[3] Nevertheless, as long as the null hypothesis $\beta^{j,k} = 0$ is rejected the term structure contains significant information about future changes in the inflation rate. In the empirical part we focus on this hypothesis.

2.2 Variation and stability of the formation content

To see why the information content varies over time, we follow Mishkin (1990), who shows that the estimated slope parameter $\beta_t^{j,k}$ in the inflation rate regression can be written as

$$\beta_t^{j,k} = \left(\sigma_t^{\,2} + \rho_t\sigma_t\right) \Big/ \left(1 + \sigma_t^{\,2} + 2\rho_t\sigma_t\right), \tag{7}$$

where $\sigma_t^{\,2} = Var\left(E_t\left(\pi_t^j - \pi_t^k\right)\right) \Big/ Var\left(E_t\left(rr_t^j - rr_t^k\right)\right)$ is the ratio of the variance of the expected inflation rate changes and the variance of the expected real rate changes, and $\rho_t = Corr\left(E_t\left(\pi_t^j - \pi_t^k\right), E_t\left(rr_t^j - rr_t^k\right)\right)$ is the correlation between the expected inflation rate changes and

3. For example. if the term premium followed an ARMA process, a better estimate could be obtained by taking this into account. However, as long as the term premium is orthogonal to the interest rate spread, the

the expected real rate changes.[4] The σ_t^2 can be thought of as signal-to-noise ratio where the signal is the information in the term structure about expected changes in inflation rates, and the noise is variation in the real-term-structure slope. Thus, the variation in the estimated slope parameter can be explained by changes in expected inflation rates, changes in expected real returns or both.

The above relations are not structural, and are likely to vary with changes in underlying factors, such as the choice of exchange rate or monetary policy regime, the degree of capital movement liberalisation and the institutional structure of bond markets. For example, the information content of yield spreads regarding future inflation may be more limited with fixed exchange rates (Gerlach and Smets, 1997). Specifically, with fixed exchange rates, movements in short-term rates aimed at defending a specific exchange rate in situations of speculative attacks may obscure the information content of yield spreads regarding future inflation to the extent that these attacks are not closely related to variations in the inflation outlook. As concerns the choice of monetary policy regime, credible strict inflation targeting as opposed to strict output targeting may lead to the term structure being uninformative about inflation. Specifically, as the short-term interest rate is set to hit the inflation target exactly (in expectations), there is no further expectations value to it, and the spread will loose any independent information about inflation, other than that contained in the target and current inflation and macro variables (Estrella, 1998). Market regulation may also matter. Specifically, more regulated financial markets may exhibit lower informational efficiency than regulated markets to the extent that market regulations induce portfolio choices of the investors that they would not have made in the absence of such regulations. These could prevent the resulting bond prices and interest rates from accurately reflecting market participants' expectations about future inflation developments. As regards the market structure, the introduction of derivative instruments could improve the informational efficiency of the prices of the underlying bonds for two reasons. First, additional traders

coefficient estimate will be consistent and have a probability limit equal to one.

4. This formula is equivalent to the standard formula for the slope parameter coefficient in a forecast equation with rational expectations, since the latter implies that the covariance of the inflation forecast error with the slope of the real term structure is equal to zero. The existence of time-varying term premia is ignored.

may be attracted. Second, to the extent that transaction costs are lower in derivatives than in spot markets, trading can take place more frequently and quickly, thus conveying additional information to the spot market participants. Similarly, the more market-driven techniques to place government bonds on the primary market are used, for example auctions instead of direct placements with a group of underwriters, the easier and faster the process of price discovery in the secondary market (De Broeck et al., 1998). All these factors vary not only across countries but also over time, which suggests the usefulness of a test for stability of the relationship between the term structure and inflation.

The previous studies which have addressed the issue of stability of the relationship (Mishkin, 1990*a*, *b*) are based on the assumption that the breakpoint is known ex ante, and thus can be treated as exogenous. However, this may not be appropriate here as one does not know the precise timing of regime changes in many cases, and even less is known about the timing of the effect of a change. Consider the dates proposed by Mishkin for monetary policy regime changes in the United States. While the choice of the first date (October 1979) coincides with the Fed's announcement of a change in its monetary operating procedures and the start of the tenure of Chairman Volcker, the timing of the effects of these changes remains uncertain.[5] Moreover, the timing of the second regime change (October 1982) is even more questionable, as the de-emphasis of monetary aggregates was a somewhat gradual process. Identifying the effects of changes in exchange rate regimes also faces similar problems. For example, the membership of the United Kingdom in the ERM may not have had an immediate effect on the information content of interest rates, as the authorities did not have to aggressively defend the exchange rate parity by raising short-term interest rates right from the beginning of ERM membership in October 1990.[6] Identification of changes in the degree of financial market regulation and the timing of their effects are also difficult.

5. On a similar note, Schwert (1986) criticises the choice of these two dates by Huizinga and Mishkin (1986) on the grounds that they are not based on direct measures of monetary policy variables, and thus remain necessarily subjective.

6. Note also that the period of ERM membership (October 1990 to September 1992) was relatively short. Thus, the distortion of the information content of interest rates may be more limited than in those countries which were subject to several episodes of speculative attacks, such as France. This may have induced

Specifically, financial market regulation involves many different dimensions which are difficult to weigh against each other, and the interactions between the different dimensions are not well-known. Moreover, the timing may be uncertain not only because of a gradual adjustment or "learning", but also because of "announcement effects", i.e. when an effect occurs in advance of the intervention. All these examples show that the assumption of ex ante knowledge of the exact timing of the breakpoint may not always be appropriate. In such situations rolling Chow tests do not offer a valid alternative, as they determine the candidate breakpoint date endogenously. However, when the breakpoint is not exogenous, conventional hypothesis testing is not valid (Banerjee, Lumsdaine and Stock, 1992).

An alternative, proposed by Andrews (1993) and Ghysels, Guay and Hall (1997), is applied here. Rather than setting the break point *a priori* and testing for its statistical significance, these techniques explicitly search for a break point. The key difference is that instead of testing for a break at a particular point in the time series, these techniques look over almost the entire sample for the location of a break point should one exist. The advantage is that the break might not be where one thinks it is, and the procedures make no *a priori* assumptions about its location. Both tests are rather general and are valid for any model estimated by generalised method of moments, such as ours which estimates the variance covariance matrix allowing for autocorrelated and heteroscedastic errors (Newey and West, 1987). Three test statistics will be used, the *supremum* Lagrange multiplier test $\sup_{t\in\tau} LM_t$ proposed by Andrews and two variations of the *supremum* predictive test $\sup_{t\in\tau} TS_t$ proposed by Ghysels, Guay and Hall. The LM test uses the estimates under the null hypothesis of no break to detect a departure from zero of the orthogonality condition in the two sub-samples, where the sample split is defined by the breakpoint date under consideration. One $\sup_{t\in\tau} TS_t$ test estimates the parameters from the first subsample up to the possible breakpoint to construct fitted values for the second sub-sample, $\sup_{t\in\tau} TS1_t$, and the other one estimates the

Gerlach and Smets (1997) to include the United Kingdom in the group of countries with floating exchange parameters from the second sub-sample to construct fitted values for the first sample, $\sup_{t\in\tau} TS2_t$. In each case departures from zero of the orthogonality conditions in the other subsample are tested. Note that, for technical reasons, these tests require excluding some fraction of the observations at the beginning and end of sample.[7] More formally, using τ to denote the set of possible breakpoints, $\tau \subset (\omega, \omega T)$, where $\omega \in (0,1)$ defines the length of that interval and T is the total number of observations, the fraction ω is required to be greater than 0. The choice of ω, following Banerjee *et. al.* called the trimming parameter, entails a trade-off between needing enough observations in the shortest regression to generate reliable parameter estimates and wanting to capture possible breakpoints early and late in the sample.[8]

It is interesting to compare the results with those obtained from the more traditional approach where a break point is assumed to be known exactly *ex ante*, and its statistical significance tested. To assess the statistical significance, the corrected critical values from Andrews (1993) are used. Essentially, the set τ collapses to one single observation, and ω (π_0 in Andrews notation) is equal to 0.50. A list of possible break dates, related to the factors referred to in the previous studies discussed above, is given in Table 3.[9]

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rates in their empirical analysis.

7. See Andrews (1993) p. 838.

8. One limitation of the technique used here is that it allows for only one break in the time series. However, this limitation is not of great importance, as the primary focus here is on whether or not the relationships are stable. In any case, the technique can easily be adapted to testing for multiple breakpoints by subsequently applying the test (Bai and Perron, 1998).

9. Various studies are drawn on. For example, the dates relating to the structure of government bond markets, e.g. the introduction of derivative instruments or new more market-driven bond placement techniques follows De Broeck *et al*., and expands their list. The identification of major capital movement liberalisation and interest rate or credit deregulation measures follows Edey and Hviding (1995), who have identified these two aspects as key dimensions of financial market deregulation.

3. Empirical analysis

The relationships between the nominal term structure and future inflation changes (6) are estimated, where *j* is varied from two to five and *k* from one to four years. As is well known the use of overlapping data induces serial correlation in the regression errors of order MA(12*j*-1), and the standard heteroscedasticity and autocorrelation-consistent Newey and West (1987) estimator could produce misleading inferences in small samples such as the present one. Therefore, empirical probability values are reported, which are calculated using a bootstrapping procedure.[10] They are in general higher than the asymptotically consistent Newey-West probability values because the empirical distribution has thicker tails, implying a lower probability of rejecting the null hypothesis (of no information content of the term structure).

Selected results are shown in Table 2. The term structures in Canada, Germany, the United States and the United Kingdom appear to be informative about future inflation in the sense that, at least for some maturity combinations, $\beta_t^{j,k}$ is positive and the null hypothesis that $\beta_t^{j,k}$ is equal to zero could be rejected at conventional levels of significance. The most informative maturity combination, in terms of maximising the R^2 , differs across countries (e.g. 5-3 years in Germany, 5-1 years in Canada, 4-1 years in the United States and 2-1 years in the United Kingdom). In contrast to the former results, there is no significant information content in the French, Italian and Japanese term structures. This confirms the results of earlier studies (e.g. Jorion and Mishkin, 1991); specifically that the information content of the term structure differs considerably across countries. In the following the description focuses on the most informative maturity combinations in the former four countries.

To see why the estimates of the slope coefficients differ, estimates of the two variables in (7), the ratio of standard deviations σ_t and the correlation coefficient ρ_t are calculated, assuming that the relevant information for the calculation of *ex ante* real interest rates can be obtained from *ex post* real rate

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10. See Schich (1999) for a detailed description of the calculations.

regressions.[11] The estimates (Figure 1) show that the more the signal-to-noise ratio exceeds one, the greater is the likelihood that the information content of the term structure is significantly different from zero. For example, in Germany, signal-to-noise ratios greater than one are always reflected in significant parameter estimates. In Canada, the United Kingdom and the United States, a signal-to-noise ratio above one is a necessary but not sufficient condition for significant parameter estimates; it has to be well above one. Note that in the case of Italy, France and Japan, the insignificant parameter estimates are reflected in signal-to-noise ratios almost always below one. Summarising, the differences in the estimated slope coefficients across countries largely reflect differences in the signal-to-noise ratio.

Do the results vary over time? Rolling regressions with a ten-year moving window demonstrate that there is also substantial variation in the parameter estimates (Figure 2) and the R^2 (Figure 3) over time. The variation in $\beta_t^{j,k}$ reflects both changes in the signal-to-noise ratio σ_t and in the correlation coefficient ρ_t. For example, in the United Kingdom, the slope coefficient estimate falls rapidly to zero as the signal-to-noise ratio drops just below one around 1987, after having varied around 1 to 1.2 during the early 1980s. Another example is the United States, where the increase of the correlation coefficient is so strong that even signal-to-noise ratios greater than one are not sufficient to guarantee significant parameter estimates. Thus, in all the four countries where the information content in the term structure is found to be significant, the results are subject to substantial changes over time, reflecting variation both in the signal-to-noise ratio and the correlation coefficient.

To see whether the results are stable, tests for a structural break with an unknown breakpoint date are applied in the case of the four countries where the term structure is found to be informative. The

11. To calculate estimated values of the two variables, the procedure outlined in Mishkin (1990b) is followed. The *ex post* real interest rate differential at each time t is regressed on the current zero-coupon yield (or yield-to-maturity) spreads and the lagged inflation rate changes to the extent that their values were available at time t. The thus fitted *ex post* real interest rate is interpreted as the expected real interest rate. This is then subtracted from the (nominal) zero-coupon yield to obtain expected inflation rates. Experiments with different specifications of the real rate regression, where the number of variables and lags were changed, show that the estimated values of the two variables are quite robust.

trimming parameter ω is set equal to 0.2, that is 20% of the total observations are not included in the set of possible breakpoint dates on each side. Figure 4 shows the test statistics for the set of possible breakpoint dates using the most informative maturity combination, and Table 4 shows the $\sup_{t\in\tau} LM_t$, $\sup_{t\in\tau} TS1_t$ and $\sup_{t\in\tau} TS2_t$ for every maturity combination. In no case is the hypothesis of structural stability rejected. Nevertheless, the figures show that the results differ between countries. In the case of Germany, all three test statistics show a very flat profile. In the case of Canada somewhat higher LM-test statistics are found only at the beginning of the interval, but this may reflect that the short regression is not efficiently estimated. Otherwise the profile is flat as well, with no indication of a structural break. By contrast, in the case of the United Kingdom and the United States the test statistics are close to the 10% critical values around the end of the 1970s and beginning of the 1980s, respectively. This raises the question how robust the results are with respect to the choice of the trimming parameter ω.

Alternative values for the trimming parameter ω are considered; specifically it is increased from 0.2 by 0.05 each time, until 0.4. We find that when the parameter is either 0.3, 0.35 or 0.4, the hypothesis of structural stability could be rejected in the case of the United Kingdom at the 10% level on the basis of the LM-test, with the breakpoint date implied by the supremum being 1981:04. Similarly, the LM-test rejects stability in the case of the United States at the 10% level of significance when the trimming parameter is either 0.35 or 0.4, with the implied breakpoint date being 1978:03. These rejections reflect the fact that the critical values for the 10% level of significance decrease as the interval is reduced. For example, the critical value is 8.57 for a trimming parameter of 0.3, 8.06 for on of 0.35 and 7.45 for one of 0.4 (as shown in Figure 4 for easy reference), while it is 9.59 for a trimming parameter of 0.2. In the cases of both the United States and the United Kingdom the slope coefficient estimates is much lower in the second than in the first sub-sample, reflecting that the term structure has become less informative. By contrast, the estimates for Germany and Canada appear to be structurally stable, and the stability results are robust to the choice of the trimming parameter.

These results are consistent with those obtained from tests where the breakpoints are assumed to be known exactly *a priori* (Table 5). While the hypothesis of stability is never rejected in the case of Germany and Canada, it is rejected in the case of the United States and the United Kingdom for several dates, mostly close to the breakpoint date implied by the tests for an unknown breakpoint. The results are broadly consistent with the results from the tests with unknown breakpoints. Specifically they suggest no structural breaks for Germany and Canada, but point to structural breaks for the United States and the United Kingdom for the late 1970s and early 1980s.

Summarising, there is substantial variation over time in the results regarding the information content, thus confirming the picture that emerges from a survey of the earlier literature. Despite this variation, the hypothesis of structural stability of the relationship between the term structure slope and future inflation could not be rejected in the cases of Germany and Canada. By contrast, in the cases of the United States and the United Kingdom, the results depend on the choice of the interval over which structural breaks are tested for. While choosing a large interval does not reveal any significant evidence of a structural break, shortening the interval yields evidence of structural breaks for both countries, as the critical values are smaller. The latter results are broadly consistent with those obtained from direct tests for a breakpoint, where the breakpoint dates are assumed to be known *a priori*.

4. Interpretation

The present section provides a tentative interpretation for the lack of a significant relationship between the term structure and future inflation changes in Italy, France and Japan and for the structural instability in that relationship in the cases of the United States and the United Kingdom. As regards the interpretation of the implied breakpoints in the cases of the United States and the United Kingdom, caution is required. Even when the coincidence of timing with any of the events identified above appears striking, this does not necessarily imply causality.[12] Notwithstanding this caveat, the following interpretations are suggested. In the case of the United States, the breakpoint date implied by our tests for an unknown breakpoint is in the late 1970s. This is consistent with the interpretation of Mishkin (1990*a*, *b*) that a structural break in the relationship between the term structure and inflation in the United States is related to changes in the Fed's monetary policy. Specifically, in the late 1970s the Fed changed its monetary operating procedures, away from nominal interest rate targeting towards the targeting of non-borrowed reserves, i.e. a component of monetary aggregates. In the early 1980s, the Fed moved away from this approach and towards a more eclectic approach to monetary policy, considering a broad range of indicators with varying weights when setting policy rates. This approach may be closer to some form of inflation targeting than the approach adopted before the late 1970s, and this may have induced the term structure to lose some of its predictive power.

In the case of the United Kingdom, the implied breakpoint date is at the beginning of the 1980s. This broadly coincides with the abolishment of foreign exchange controls in the late 1970s and interest rates deregulation in the early 1980s. However, while these measures would be expected to be associated with an increase in the information content, measured in terms of the coefficient of determination the information content actually decreased after the structural break. The beginning of the 1980s was also characterised by increasing uncertainty about the usefulness of the intermediate monetary aggregates targets for monetary policy. This led to more emphasis being placed on exchange rate movements as nominal anchors. The stronger focus on exchange rate targets during the second half of the 1980s may have implied that the setting of short-term interest rates became less closely linked to domestic developments, with the effect that the term structure became less informative about domestic inflation, than during the 1970s.

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12. Consider the example of a monetary policy regime change. A formal test for causality would require a more precise definition of a "regime changes", with a view to characterising "regimes" by specific functional relationship between relevant economic variables on the one hand, and monetary policy variables on the other.

As regards the variation of results across countries, our results are consistent with two explanations proposed in the earlier literature. First, the choice of exchange rate regime may be partly responsible for the variation of results across countries. France and Italy which effectively pegged their exchange rates to the D-Mark, underwent several episodes of exchange rate attacks (see also Gerlach and Smets, 1997), possibly distorting the information content of the term structure regarding the domestic inflation outlook. By contrast, in Canada, Germany, the United States and the United Kingdom, exchange rates have been relatively freely floating during the sample period, at least most of the time. However, while the absence of episodes of speculative attacks under pegged exchange rates may be a necessary condition for a significant information content, the example of Japan suggests that it is not sufficient. Second, the variation of results across countries may reflect the degree of deregulation of financial markets. The picture that emerges from an account of major deregulation measures is one of two groups of countries. On the one hand, capital controls and interest rate or credit regulations have been abolished before or around the early 1980s in Canada, Germany, the United Kingdom and the United States -- countries where the term structures are found to be informative. On the other hand, some of the capital controls or interest rate regulations were only abolished during the 1990s in France, Italy and Japan[13] -- countries where the term structures are not found to be informative. Notwithstanding the fact that the above considered aspects do not provide an exhaustive description of the degree of financial market regulation, our empirical results indeed lend some weak support to the hypothesis that the variation of information-content results across countries reflects differences with respect to the deregulation of financial markets.

13. For example, in Japan the "authorised foreign exchange bank" system was eliminated and the prior notification requirement for capital transactions abolished only in April 1998. The regulation of interest rates for small denomination deposits was abolished in July 1993. In France, the Article 67 of the Treaty of Rome (1988 -- liberalisation of capital movements) was fully implemented in January 1990. While quantitative control on credit was abandoned in 1987, banks were allowed to freely set interest rates on time deposits and bank bills of at least one month in January 1990. While the ceiling on bank lending in Italy expired in March 1988, the Bank of Italy continued to use "moral suasion" to ensure that credit growth remained within set limits. The Article 67 was fully implemented in May 1990. For details of liberalisation and deregulation measures see Takeda and Turner (1992), Passacantando (1996) and the references given in Table 3.

5. Conclusions

One conclusion is that there is significant information content in the term structure of interest rates of the United States, Germany, Canada, and the United Kingdom. This is an important result because previous empirical studies have provided mixed evidence for some of the countries for which we find the term structure to be informative. Another conclusion is that the information content of term structures about future inflation changes is subject to variation not only across countries but also over time, confirming the picture that emerges from a survey of the earlier literature. Nevertheless, the hypothesis of structural stability of the relationship between the term structure slope and future inflation could not be rejected in the cases of Canada and Germany. By contrast, there is evidence for structural breaks in the cases of the United States and the United Kingdom. These possibly reflect changes in the monetary policy regimes. Changes in financial market deregulation measures or changes in the institutional structure of bond markets did not seem to have had a significant impact on the information content of the term structures of interest rates. Thus, from a monetary policy point of view, the following conclusion can be drawn. The term structure of interest rates is often a good indicator of future inflation developments. However, its usefulness may change over time, in particular with policy regime changes, and should therefore be regularly assessed on the basis of empirical evidence.

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Table 1. **Selected tests of the information content of the term structure regarding inflation**

Canada

Day and Lange (1997)

Sample	Maturity combination	Significance[1]
1967:01 - 1990:02	5 - 1	(*)
1967:01 - 1991:02	4 - 1	(*)
1967:01 - 1991:02	3 - 1	
1967:01 - 1993:02	2- 1	

Germany

Gerlach (1997)

Sample	Maturity combination	Significance[1]
1968:01 - 1993:01	2 - 1	
1968:01 - 1992:01	3 - 1	
1968:01 - 1991:01	4 - 1	
1968:01 - 1990:01	5 - 1	*
1968:01 - 1992:01	3 - 2	
1968:01 - 1991:01	4 - 2	**
1968:01 - 1990:01	5 - 2	**

Schich (1999)

Sample	Maturity combination	Significance[1]
1972:09 - 1995:10	2 - 1	
1972:09 - 1994:10	3 - 1	
1972:09 - 1993:10	4 - 1	**
1972:09 - 1992:10	5 - 1	***
1972:09 - 1994:10	3 - 2	*
1972:09 - 1993:10	4 - 2	**
1972:09 - 1992:10	5 - 2	**

Jorion and Mishkin (1991)

Sample	Maturity combination	Significance[1]
1973:08 - 1987:06	2 - 1	
1973:08 - 1986:06	3 - 1	
1973:08 - 1985:06	4 - 1	*
1973:08 - 1984:06	5 - 1	**

Koedijk and Kool (1995)

Sample	Maturity combination	Significance[1]
1976:04 - 1990:09	2 - 1	
1976:04 - 1989:09	3 - 1	
1976:04 - 1988:09	4 - 1	(*)
1976:04 - 1987:09	5 - 1	(*)
1982:01 - 1990:09	2 - 1	
1982:01 - 1989:09	3 - 1	
1982:01 - 1988:09	4 - 1	
1982:01 - 1987:09	5 - 1	

United States

Jorion and Mishkin (1991)

Sample	Maturity combination	Significance[1]
1973:08 - 1987:06	2 - 1	**
1973:08 - 1986:06	3 - 1	**
1973:08 - 1985:06	4 - 1	*
1973:08 - 1984:06	5 - 1	

Koedijk and Kool (1995)

Sample	Maturity combination	Significance[1]
1976:04 - 1990:09	2 - 1	
1976:04 - 1989:09	3 - 1	
1976:04 - 1989:09	4 - 1	
1976:04 - 1987:09	5 - 1	
1982:01 - 1990:09	2 - 1	
1982:01 - 1989:09	3 - 1	
1982:01 - 1988:09	4 - 1	
1982:01 - 1987:09	5 - 1	

Table 1 (continued). **Selected tests of the information content of the term structure regarding inflation**

United Kingdom

Jorion and Mishkin (1991)			Breedon and Chadha (1996)		
Sample	Maturity combination	Significance[1]	Sample	Maturity combination	Significance[1]
1973:08 - 1987:06	2 - 1	**	1982:01 - 1994:03	2 - 1	
1973:08 - 1986:06	3 - 1	*	1982:01 - 1993:03	3 - 1	
1973:08 - 1985:06	4 - 1	*	1982:01 - 1992:03	4 - 1	
1973:08 - 1984:06	5 - 1		1982:01 - 1991:03	5 - 1	
			1982:01 - 1993:03	3 - 2	
			1982:01 - 1992:03	4 - 2	
			1982:01 - 1991:03	5 - 2	

France

Jondeau and Ricart (1997)			Koedijk and Kool (1995)		
Sample	Maturity combination	Significance[1]	Sample	Maturity combination	Significance[1]
1980:01 - 1993:12	2 - 1		1982:01 - 1990:09	2 - 1	
1980:01 - 1992:12	3 - 1		1982:01 - 1989:09	3 - 1	
1980:01 - 1991:12	4 - 1		1982:01 - 1988:09	4 - 1	
1980:01 - 1990:12	5 - 1		1982:01 - 1987:09	5 - 1	
1985:01 - 1993:12	2 - 1	*			
1985:01 - 1993:12	3 - 1	*			
1985:01 - 1993:12	4 - 1	***			
1985:01 - 1993:12	5 - 1	*			

Japan

Koedijk and Kool (1995)		
Sample	Maturity combination	Significance[1]
1982:01 - 1990:09	2 - 1	
1982:01 - 1989:09	3 - 1	
1982:01 - 1988:09	4 - 1	
1982:01 - 1987:09	5 - 1	

1. Significance level at which the hypothesis that the slope parameter in a regression of forward inflation rate changes on the spread is equal to zero could be rejected. Significance at the 1% level is denoted by ***, at the 5% level by **, and at the 10% by *, significance levels being obtained via bootshapping or Monte Carlo simulations. Where only standard Newey-West (asymptotically correct) probability values reported, a (*) signifies significance at least at the 10% level. Note that in the latter cases the use of asymptotically correct, as opposed to empirical, probability values tends to lead to overrejection of the null hypothesis.

Table 2. **Regressions of forward inflation changes on yield spreads**

Spread (j-k)	Period	Inflation regressions α	β	R^2
		United States		
2-1	1961:06 - 96:02	-0.25 [0.02]	1.26 [0.00]**	0.19
3-1	1961:06 - 95:02	-0.49 [0.02]	1.47 [0.01]**	0.28
4-1	1961:06 - 94:02	-0.61 [0.06]	1.45 [0.03]*	0.30
5-1	1961:06 - 93:02	-0.62 [0.13]	1.38 [0.07]	0.30
3-2	1961:06 - 95:02	-0.25 [0.01]	1.68 [0.02]*	0.25
4-2	1961:06 - 94:02	-0.34 [0.12]	1.44 [0.08]	0.20
5-2	1961:06 - 93:02	-0.34 [0.30]	1.27 [0.16]	0.16
4-3	1961:06 - 94:02	-0.08 [0.48]	0.89 [0.24]	0.06
5-3	1961:06 - 93:02	-0.08 [0.70]	0.69 [0.44]	0.03
5-4	1960:06 - 93:02	0.01 [0.92]	0.06 [0.96]	0.00
		Japan		
2-1	1980:03 - 96:05	-0.10 [0.19]	0.72 [0.16]	0.04
3-1	1980:03 - 95:05	-0.23 [0.32]	0.56 [0.58]	0.03
4-1	1980:03 - 94:05	-0.33 [0.36]	0.13 [0.88]	0.00
5-1	1980:03 - 93:05	-0.45 [0.37]	0.14 [0.85]	0.00
3-2	1980:03 - 95:05	-0.06 [0.62]	0.07 [0.94]	0.00
4-2	1980:03 - 94:05	-0.15 [0.54]	-0.04 [0.94]	0.00
5-2	1980:03 - 93:05	-0.25 [0.54]	-0.12 [0.84]	0.01
4-3	1980:03 - 94:05	-0.05 [0.62]	0.00 [0.99]	0.00
5-3	1980:03 - 93:05	-0.12 [0.63]	-0.20 [0.69]	0.02
5-4	1980:03 - 93:05	-0.04 [0.74]	-0.18 [0.63]	0.02
		Germany		
2-1	1967:01 - 96:01	-0.10 [0.38]	0.26 [0.29]	0.03
3-1	1967:01 - 95:01	-2.28 [0.24]	0.46 [0.13]	0.09
4-1	1967:01 - 94:01	-0.49 [0.17]	0.66 [0.05]*	0.15
5-1	1967:01 - 93:01	-0.70 [0.12]	0.82 [0.02]*	0.22
3-2	1967:01 - 95:01	-0.18 [0.10]	0.84 [0.04]*	0.14
4-2	1967:01 - 94:01	-0.39 [0.09]	1.10 [0.01]**	0.22
5-2	1967:01 - 93:01	-0.61 [0.06]	1.30 [0.01]**	0.29
4-3	1967:01 - 94:01	-0.21 [0.04]*	1.48 [0.00]**	0.26
5-3	1967:01 - 93:01	-0.43 [0.04]*	1.69 [0.00]**	0.34
5-4	1967:01 - 93:01	-0.21 [0.02]*	1.93 [0.00]**	0.34
		France		
2-1	1980:01 - 96:06	-0.33 [0.00]**	-0.22 [0.62]	0.02
3-1	1980:01 - 95:06	-0.65 [0.03]*	-0.26 [0.60]	0.03
4-1	1980:01 - 94:06	-0.93 [0.15]	-0.33 [0.53]	0.04
5-1	1980:01 - 93:06	-1.27 [0.21]	-0.29 [0.64]	0.02
3-2	1980:01 - 95:06	-0.32 [0.05]*	-0.35 [0.60]	0.03
4-2	1980:01 - 94:06	-0.55 [0.15]	-0.47 [0.50]	0.05
5-2	1980:01 - 93:06	-0.84 [0.25]	-0.44 [0.53]	0.04
4-3	1980:01 - 94:06	-0.25 [0.20]	-0.72 [0.20]	0.08
5-3	1980:01 - 93:06	-0.51 [0.26]	-0.56 [0.37]	0.05
5-4	1980:01 - 93:06	-0.27 [0.22]	-0.28 [0.68]	0.01
		Italy		
2-1	1985:10 - 96:01	-0.02 [0.85]	-0.32 [0.53]	0.06
3-1	1985:10 - 95:01	0.07 [0.74]	-0.50 [0.26]	0.15
4-1	1985:10 - 94:01	-0.01 [0.97]	-0.20 [0.53]	0.04
5-1	1985:10 - 93:01	-0.20 [0.77]	-0.09 [0.79]	0.01
3-2	1985:10 - 95:01	-0.01 [0.90]	-0.71 [0.19]	0.12
4-2	1985:10 - 94:01	-0.10 [0.64]	-0.29 [0.55]	0.04
5-2	1985:10 - 93:01	-0.21 [0.57]	-0.07 [0.82]	0.00
4-3	1985:10 - 94:01	-0.13 [0.28]	-0.04 [0.90]	0.00
5-3	1985:10 - 93:01	-0.26 [0.31]	0.18 [0.61]	0.03
5-4	1985:10 - 93:01	-0.18 [0.17]	0.37 [0.51]	0.11
		United Kingdom		
2-1	1970:01 - 96.02	-0.25 [0.05]*	1.37 [0.01]**	0.14
3-1	1970:01 - 95:02	-0.53 [0.06]	1.27 [0.02]*	0.13
4-1	1970:01 - 94:02	-0.72 [0.11]	1.19 [0.05]*	0.11
5-1	1970:01 - 93:02	-0.89 [0.14]	0.87 [0.12]	0.07
3-2	1970:01 - 95:02	-0.26 [0.13]	1.07 [0.14]	0.06
4-2	1970:01 - 94:02	-0.43 [0.18]	0.82 [0.28]	0.04
5-2	1970:01 - 93:02	-0.56 [0.26]	0.45 [0.47]	0.01
4-3	1970:01 - 94:02	-0.17 [0.34]	0.45 [0.51]	0.01
5-3	1970:01 - 93:02	-0.31 [0.37]	0.10 [0.87]	0.00
5-4	1970:01 - 93:02	-0.15 [0.34]	-0.15 [0.81]	0.00
		Canada		
2-1	1967:01 - 96:05	-0.13 [0.07]	1.00 [0.04]*	0.12
3-1	1967:01 - 95:05	-0.33 [0.04]*	1.51 [0.00]**	0.27
4-1	1967:01 - 94:05	-0.53 [0.02]*	1.72 [0.00]**	0.34
5-1	1967:01 - 93:05	-0.71 [0.04]*	1.78 [0.00]**	0.36
3-2	1967:01 - 95:05	-0.21 [0.00]**	1.81 [0.00]**	0.27
4-2	1967:01 - 94:05	-0.41 [0.02]*	1.79 [0.01]**	0.29
5-2	1967:01 - 93:05	-0.58 [0.05]	1.78 [0.01]**	0.28
4-3	1967:01 - 94:05	-0.18 [0.07]	1.48 [0.05]*	0.17
5-3	1967:01 - 93:05	-0.31 [0.22]	1.30 [0.16]	0.14
5-4	1967:01 - 93:05	-0.11 [0.43]	0.78 [0.38]	0.05

Note: Empirical probability values in square brackets, obtained using a bootstrapping method with 1000 artificial samples. Rejection of the null hypothesis at the 5 per cent significance level is denoted by * and at the 1 per cent level by **.

Table 3. **Selected changes in the structure of the economy**

Country	Date	Monetary policy or exchange rate regime changes
United States	1979:10	Fed announces change in monetary policy operating procedures, start of Chairman Volcker.
	1982:10	Change from some type of money supply target to more eclectic approach.
Germany	1974:12	Bundesbank announces the adoption of its intermediate targeting strategy.
United Kingdom	1990:10	Beginning of ERM membership.
	1992:10	End of ERM membership.

Country	Date	Liberalisation of capital movements
United States	1974:01	Controls on capital outflows with the interest equalization tax (1963) were removed.
Germany	1981:03/ 1984:08	The last (and hardly significant) quantitative control, the authorisation requirement for the purchase of domestic securities with maturities of up to one year by non-residents was removed in 1981:03, but the 25 per cent withholding tax on fixed-rate securities held by non-residents was abolished only in 1984:08.
Canada		Never introduced all-encompassing controls.
United Kingdom	1979:10	Removal of foreign exchange controls.

Country	Date	Interest rate or direct credit deregulation
United States	1982:12	The Depository Institutions Deregulation and Monetary Control Act of March 1980 and the Garn-St. Germain Depository Institution Act of December 1982 eliminated deposit interest-rate ceilings.
Germany	1967:04	Abolition of regulation specifying upper limits for interest on deposits and interest paid.
Canada		No major formal interest rate or credit regulation, at least since the beginning of the 1970s.
United Kingdom	1981:06	Abolition of the "supplementary special deposit scheme" whereby banks recording faster rates of growth of deposits than had been authorised were penalised by placing non-interest bearing deposits with the Bank of England at progressively higher rates.

Country	Date	Recent structural changes in selected bond markets
Germany	1988:09/ 1989:04/ 1990:07	Introduction of Bund futures contract on LIFFE. Options on Bund futures launched on LIFFE. New issuing technique: traditional underwriting procedure combined with auction.
Canada	1980:09/ 1989:09	Introduction of futures on Treasury bonds on Toronto Stock Exchange. Introduction of government bond futures on Montreal Exchange.
United Kingdom	1986:03/ 1986:10	Introduction of options on Gilts. Tender and tap system combined with auctions; overall opening up of market (Big Bang).

Sources: Various issues of *International Capital Markets - Developments and Prospects* (IMF) and *Financial Market Trends* and *Economic Surveys* (OECD); OECD, *Banks Under Stress*, (1992); Deutsche Bundesbank *Monthly Bulletin*: March 1965, October 1967 and July 1985; Edey and Hviding (1995); Broeck and Guillaume, and Van der Stichele (1998).

Table 4. **Tests for single structural break with unknown breakpoint date**

(trimming parameter equal to 0.2)

		Canada			**Germany**			**United Kingdom**				**United States**			
Spreads	Sample	LM-test	TS1-test	TS2-test	LM-test	TS1-test	TS2-test	Sample	LM-test	TS1-test	TS2-test	Sample	LM-test	TS1-test	TS2-test
2 - 1	73:01	5.17	4.11	4.60	3.13	6.02	4.50	75:06	8.89	5.75	4.39	68:07	8.19	5.73	5.76
	- 90:12	(73:04)	(73:04)	(89:12)	(73:01)	(73:01)	(90.09)	- 91:07	(81:04)	(78:06)	(91:07)	- 89:11	(78:12)	(79:01)	(88:11)
3 - 1	72:11	5.47	4.38	3.32	2.75	4.44	2.42	75:03	5.27	4.75	2.33	68:05	8.56	5.06	5.39
	- 90:02	(82:03)	(72:11)	(89:12)	(72:11)	(72:11)	(80:02)	- 90:10	(81:04)	(75:04)	(80:05)	- 89:01	(78:12)	(79:02)	(89:01)
4 - 1	72:08	5.39	3.51	2.82	3.01	3.62	2.36	75:01	2.76	2.83	1.41	68:03	8.43	4.80	5.00
	- 89:05	(72:08)	(72:08)	(88:04)	(72:08)	(72:08)	(80:02)	- 89:12	(81:04)	(80:12)	(80:12)	- 88:03	(78:03)	(72:06)	(88:03)
5 - 1	72:06	7.18	3.63	2.58	3.80	3.29	2.33	74:11	1.61	1.67	1.91	67:12	7.66	3.90	4.55
	- 88:07	(72:06)	(72:06)	(82:04)	(79:12)	(72:06)	(80:05)	- 89:02	(74:11)	(85:02)	(88:05)	- 87:06	(78:01)	(72:07)	(87:06)
3 - 2	72:11	5.47	4.09	2.67	2.76	3.90	2.49	75:03	1.23	1.54	1.92	68:05	8.15	5.02	5.19
	- 90:02	(82:01)	(72:11)	(88:12)	(72:11)	(72:11)	(90:02)	- 90:10	(81:02)	(81:02)	(90:10)	- 89:01	(78:01)	(68:06)	(88:07)
4 - 2	72:08	4.56	2.94	2.08	4.05	3.40	1.79	75:01	1.34	1.13	1.04	68:03	8.49	4.67	4.98
	- 89:05	(72:08)	(72:08)	(78:12)	(72:09)	(72:08)	(76:06)	- 89:12	(88:02)	(78:03)	(88:05)	- 88:03	(77:11)	(71:12)	(88:02)
5 - 2	72:06	6.56	3.60	1.98	4.84	3.17	1.86	74:11	3.11	1.88	2.61	67:12	8.03	3.71	4.68
	- 88:07	(72:06)	(72:06)	(78:11)	(72:09)	(72:06)	(79:04)	- 89:02	(87:12)	(84:05)	(88:04)	- 87:06	(77:06)	(72:02)	(87:06)
4 - 3	72:08	3.68	3.66	2.44	6.16	3.38	3.04	75:01	2.75	1.60	2.58	68:03	7.98	4.07	4.98
	- 89:05	(78:07)	(72:08)	(77:02)	(72:09)	(73:04)	(89:04)	- 89:12	(87:04)	(84:07)	(87:12)	- 88:03	(77:02)	(71:12)	(87:03)
5 - 3	72:06	4.77	3.83	2.05	5.35	3.06	1.95	74:11	3.97	2.21	3.29	67:12	8.11	3.68	4.24
	- 88:07	(72:06)	(72:06)	(77:01)	(72:09)	(72:06)	(79:06)	- 89:02	(87:03)	(83:07)	(88:04)	- 87:06	(76:12)	(71:08)	(86:11)
5 - 4	72:06	5.41	3.57	2.07	3.02	2.62	2.67	74:11	2.77	1.78	3.05	67:12	6.24	3.98	2.89
	- 88:07	(77:07)	(72:06)	(77:07)	(87:09)	(72:06)	(87:09)	- 89:02	(75:12)	(83:04)	(88:05)	- 87:06	(76:10)	(68:02)	(76:01)

Explanation: LM-test and TS-test and TS1- and TS2-test are the Lagrange multiplier and "predictive failure" test (Guysel and Hall, 1990) for one break point, respectively. TS1 uses the parameter estimates from the first and TS2 from the second sample for prediction. The first row shows the value of the supremum of the test and the second row the implied breakpoint date. The asterics * and ** denote rejection of the null hypothesis of stability at the 10 and 5 per cent significance level, respectively. Critical values are obtained from Andrews (1997), and are 9.59 and 11.26 at the 10 per cent and 5 per cent level, respectively.

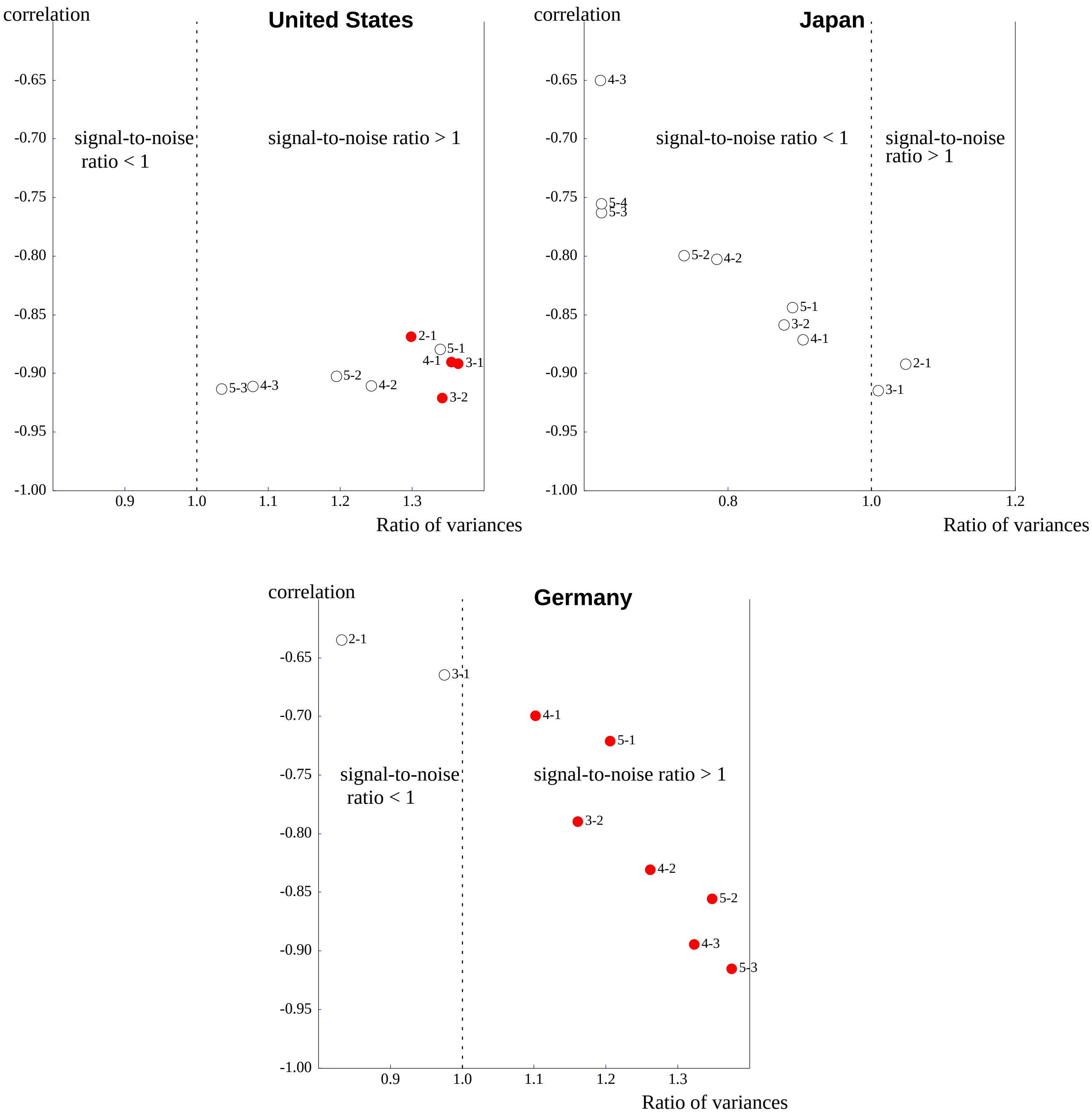
Table 5. **Tests for single structural break with known breakpoint date**

(choice of dates following the list in Table 3)

Country	Spread	Date	LM	TS1	TS2
Germany	5 - 3	1974:12	1.55	1.07	0.78
			[0.46]	[0.59]	[0.68]
		1981:03	0.51	0.45	0.59
			[0.77]	[0.80]	[0.74]
		1984:08	1.27	0.93	1.62
			[0.53]	[0.63]	[0.45]
		1988:09	0.56	1.14	2.51
			[0.75]	[0.57]	[0.29]
		1989:04	0.36	1.16	2.67
			[0.84]	[0.56]	[0.26]
		1990:07	0.31	1.22	2.36
			[0.85]	[0.57]	[0.31]
Canada	5 - 1	1980:09	2.31	1.97	1.72
			[0.32]	[0.37]	[0.42]
		1989:09	1.56	0.98	2.08
			[0.46]	[0.61]	[0.35]
United States	4 - 1	1974:01	4.21	4.06	2.54
			[0.12]	[0.13]	[0.28]
		1979:10	6.85	3.62	3.89
			[0.03]**	[0.16]	[0.14]
		1982:10	6.26	3.17	3.85
			[0.04]**	[0.20]	[0.15]
		1982:12	5.99	3.05	3.81
			[0.05]**	[0.22]	[0.15]
United Kingdom	2 - 1	1979:10	7.35	5.55	3.26
			[0.03]**	[0.06]*	[0.20]
		1981:06	8.75	5.44	3.46
			[0.01]**	[0.07]*	[0.18]
		1986:03	3.82	3.09	2.19
			[0.15]	[0.21]	[0.33]
		1986:10	3.28	2.79	1.97
			[0.20]	[0.25]	[0.37]
		1990:10	1.24	1.26	2.07
			[0.54]	[0.53]	[0.36]
		1992:09	1.58	1.22	6.16
			[0.45]	[0.54]	[0.05]**

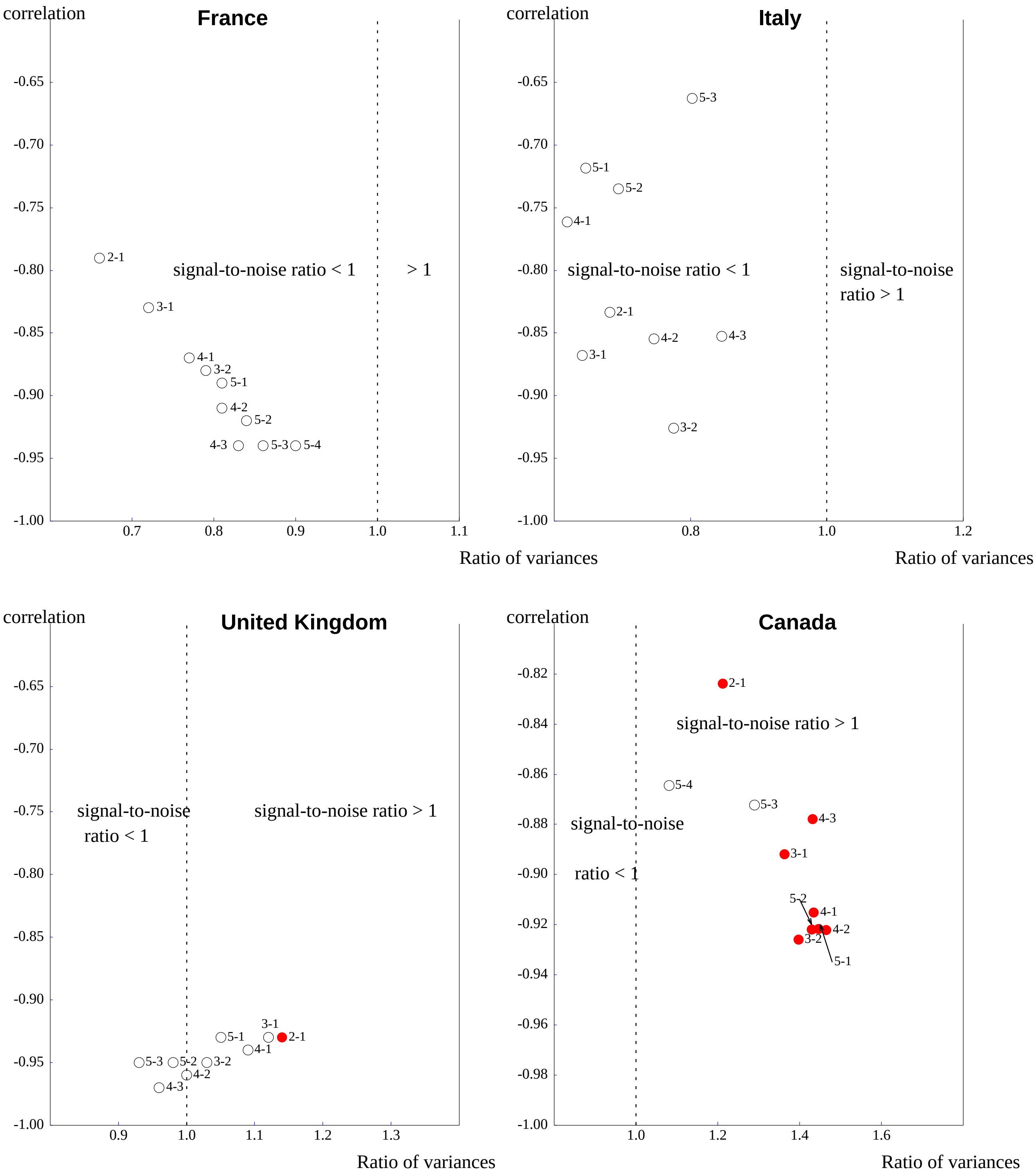
Explanation: LM-test and TS1- and TS2-tests are the Lagrange multiplier and “predictive failure” tests (Guysel and Hall, 1990) for one break point, respectively. TS1 uses the parameter estimates from the first and TS2 from the second sample for prediction. Probability values are in square brackets. The asterics * and ** denote rejection of the null hypothesis of stability at the 10 and 5 per cent significance level, respectively. Critical values are obtained from Andrews (1993).

Figure 1. Variance and correlation of expected inflation and real rate changes



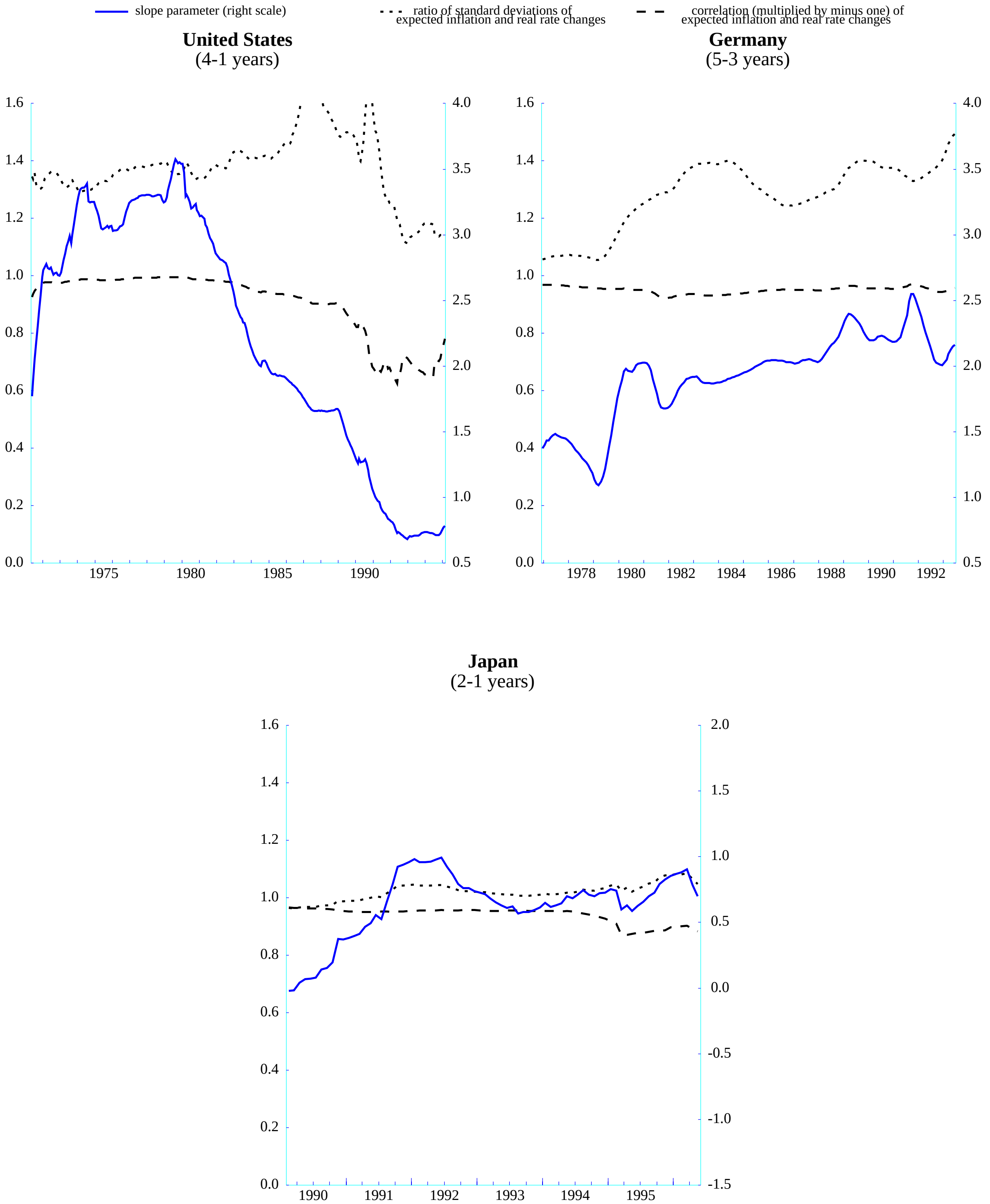
Explanation : The black dots symbolise the instances in which the slope parameter is significantly different from zero. The dotted line indicates a signal-to-noise ratio equal to one for easy reference.

Figure 1 (cont.) Variance and correlation of expected inflation and real rate changes



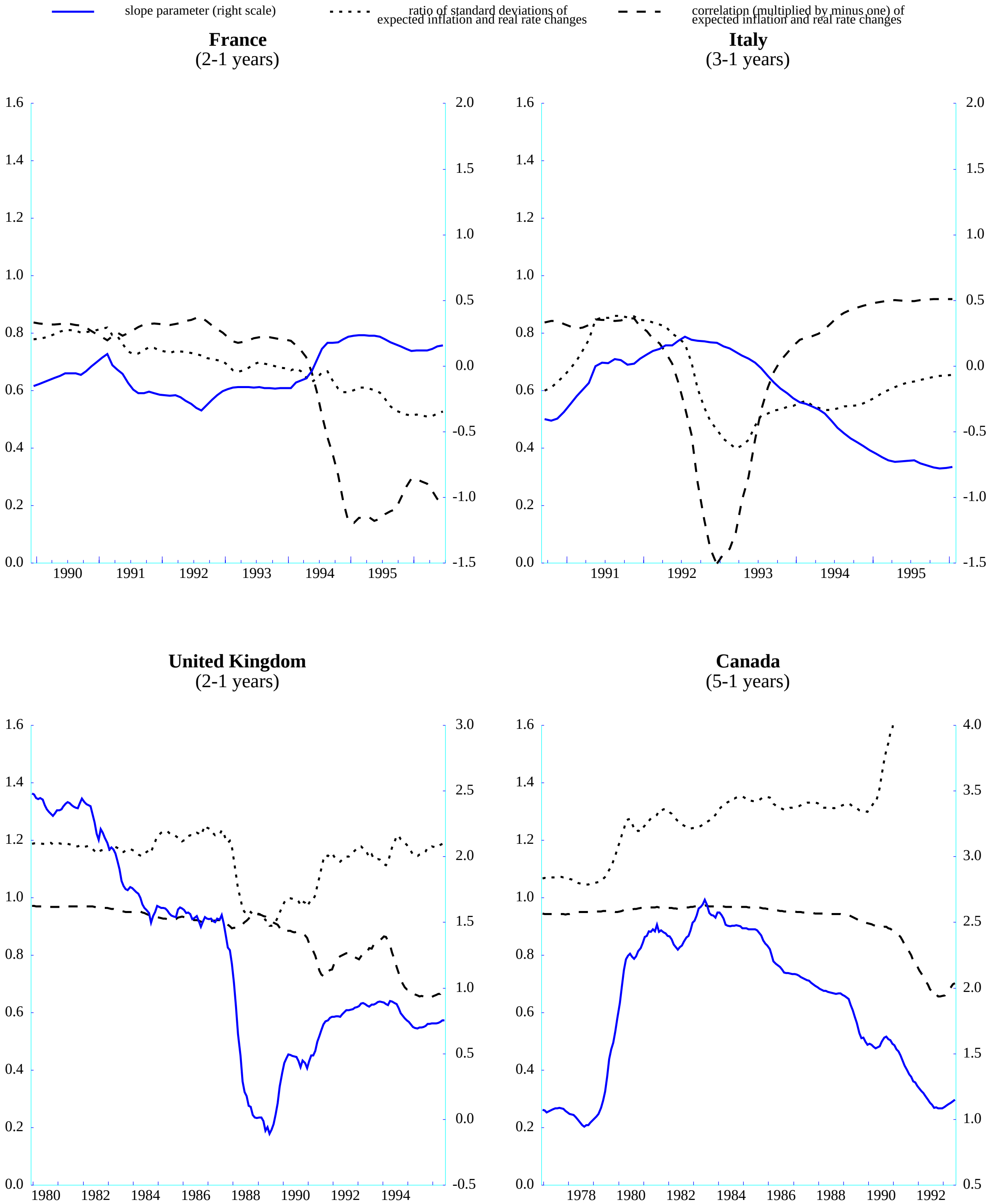
Explanation : The black dots symbolise the instances in which the slope parameter is significantly different from zero. The dotted line indicates a signal-to-noise ratio equal to one for easy reference.

Figure 2. Variances and correlation of real and inflation expectation components
(rolling 10-year window)



Note: The estimation of the three parameters refers to the end of the 10-year moving window in rolling regressions. Thus, for example, the parameter estimates shown for December 1994 refer to a regression using a sample from December 1984 to December 1994.

Figure 2 (cont.). Variances and correlation of real and inflation expectation components
(rolling 10-year window, 5-year window for Italy)

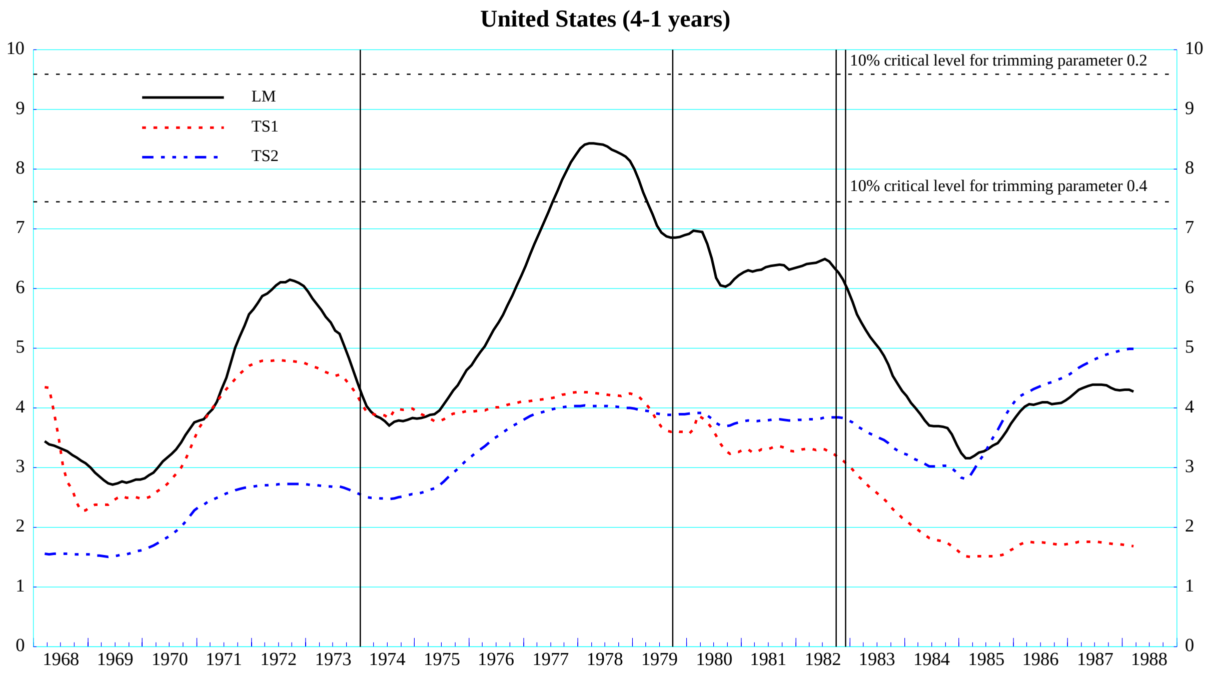


Note: The estimation of the three parameters refers to the end of the 10-year moving window in rolling regressions. Thus, for example, the parameter estimates shown for December 1994 refer to a regression using a sample from December 1984 to December 1994.

Figure 3. Change in explanatory power: R^2 from 10-year-window rolling regressions

Note: The estimated R^2 refers to the end of the 10-year moving window in rolling regressions. Thus, for example, the R^2 shown for December 1994 refers to a regression using a sample from December 1984 to December 1994.

Figure 4. Test for structural break with unknown breakpoint date



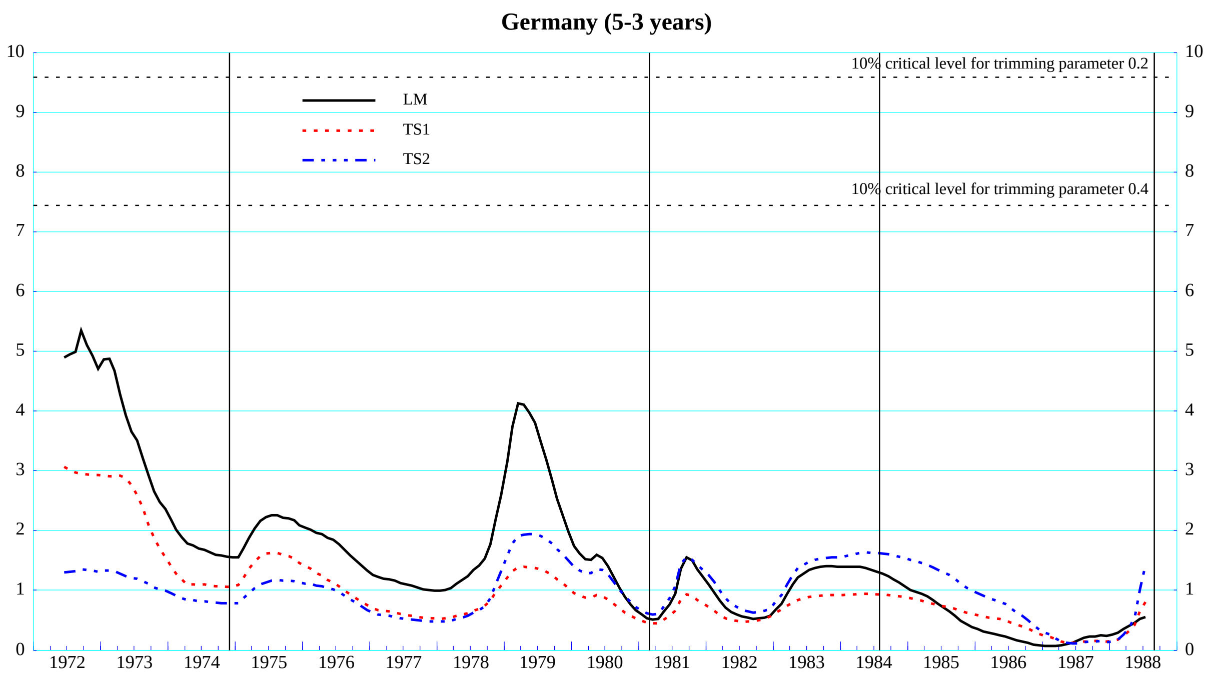


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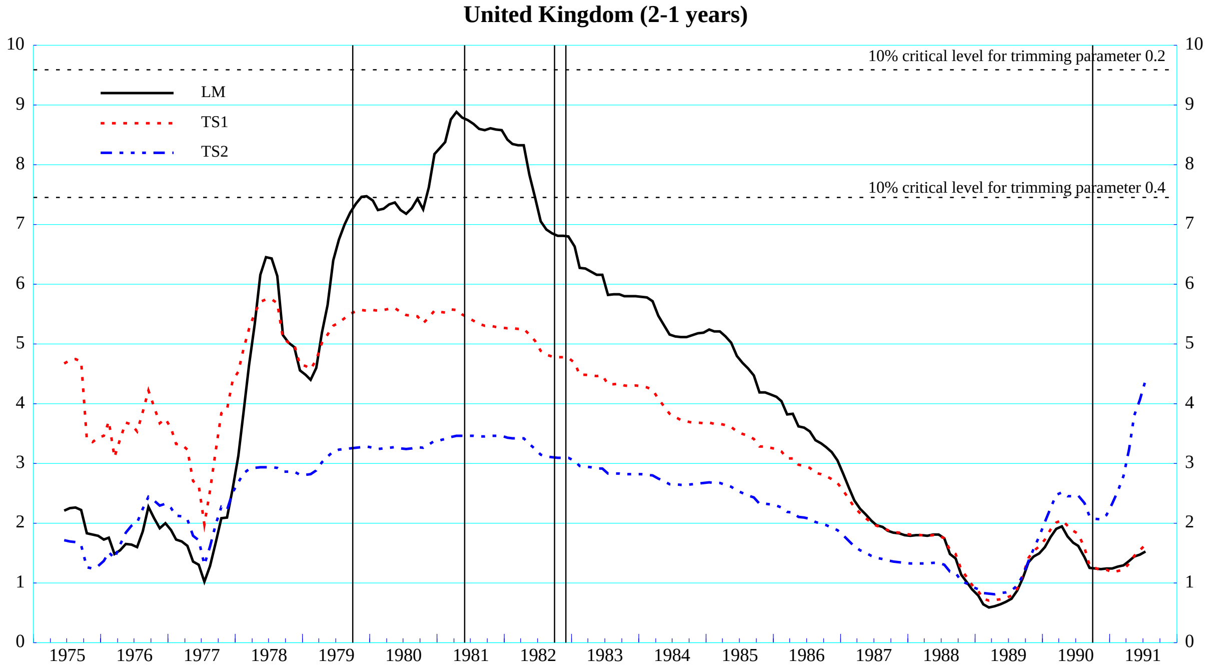
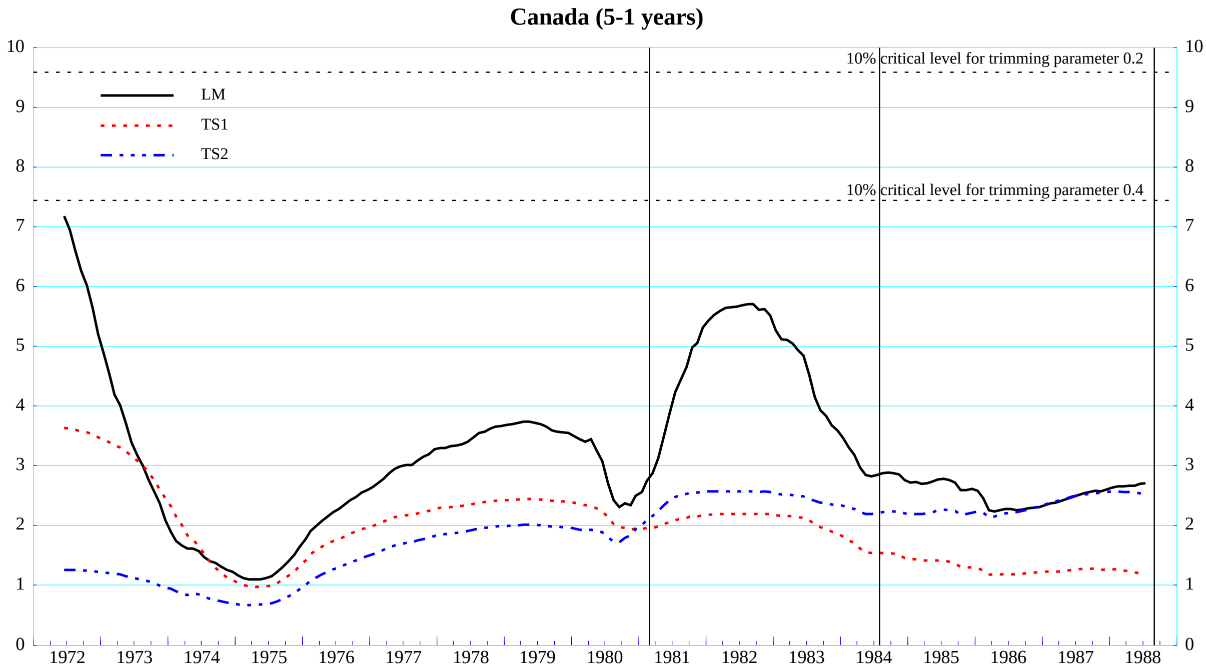
Figure 4 (cont.) Test for structural break with unknown breakpoint date





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