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FINANCIAL MARKETS' ASSESSMENT OF EMU

David S. Bates

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ABSTRACT

This article reviews the assumptions and methodologies underlying "EMU probability calculators," which infer from financial data the probability of specific countries joining the European Monetary Union. Some historical evidence is presented in support of the expectations hypothesis for intra-European interest rate differentials underlying most calculators, while various potential biases are deemed negligible. The various EMU calculators differ primarily in their scenarios for intra-European interest rate differentials conditional upon EMU not occurring. This article also discusses what can be inferred from financial data regarding future policies of the European Central Bank.

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On May 2, 1998, the leaders of the countries of the European Union formally selected those countries deemed eligible to participate in a common currency union. This union, commonly referred to as the European Monetary Union (EMU)[1], was the third and final stage of a process established under the Maastricht treaty of December, 1991. While various criteria were established in that treaty as pre-conditions for membership, the most severe constraints were with respect to the "convergence" criteria regarding acceptable levels of inflation rates, interest rates, budget deficits and national debt. An additional criterion of exchange rate stability became obsolete after the European currency crises of 1992 and 1993. 11 countries[2] were selected as eligible; an additional 3 (Denmark, Sweden, and the U.K.) chose not to participate, while Greece failed to meet the convergence criteria.

Although the EMU now appears virtually certain to begin on January 1, 1999 with broad participation, that outcome was not expected over most of the 1992-98 period. Academics discussed a two-tiered Europe consisting of a few core countries that would meet the convergence criteria, and a broader group that would not qualify until later. Paralleling this uncertainty, various academic researchers and banks developed "EMU probability calculators" to infer from financial data the probability of various countries qualifying for admission. The best-known is probably the one developed by J.P. Morgan, the results of which were regularly reported in the *Financial Times.* Four such probability calculations for Italy are illustrated below in Figure 1.[3]

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[1]Technically, EMU is the acronym used by the European Commission for Economic and Monetary Union, of which the currency union is the final stage.

[2]Austria, Belgium, Finland France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal, and Spain.

[3]The Favero *et al* (1997) and Lund (1998) EMU probability assessments are the quarterly and weekly numbers, respectively, provided by the authors. The De Grauwe (1996) and Morgan (1997) estimates are representative monthly rates estimated from published graphs of daily data.

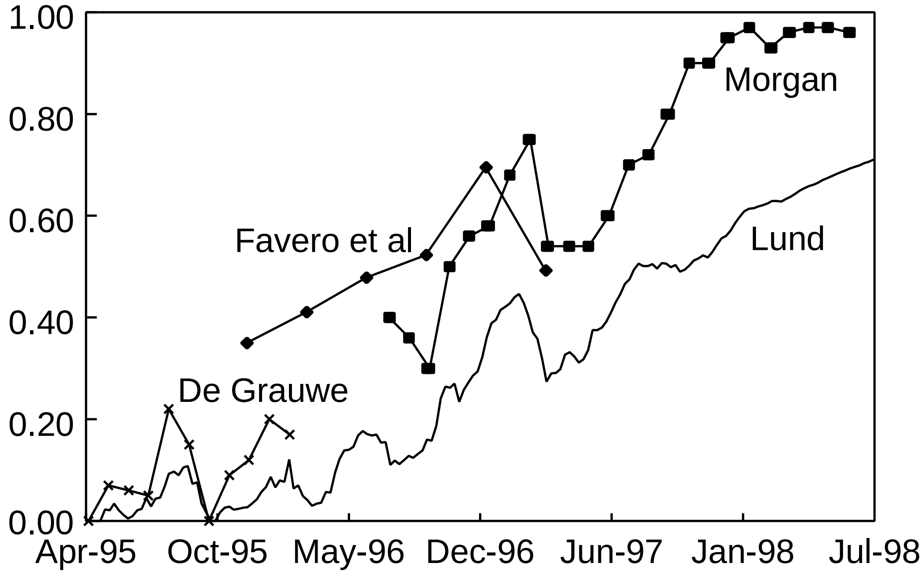

Figure 1. EMU probability assessments for Italy, 1995-98. Lund's (1998) and Favero *et al*'s (1997) estimates are inferred probabilities of Italy joining the EMU during 1999-2001. The time frame for the other two estimates is roughly comparable.

This article primarily reviews the methodologies employed in constructing such calculators. While the initial group of participants has now been determined, many other countries are waiting in the wings and may join later. A retrospective examination of the methodologies may therefore be useful in providing a guide to constructing new calculators for prospective entrants. Section I discusses two alternative direct approaches (Arrow-Debreu contracts, options-based assessments) to creating EMU probability assessments, while Section II examines the more common approaches that use the term structures of European interest rates. Section III concludes with some discussion of what can be inferred from financial data regarding future policies of the new European Central Bank.

I. Non-standard EMU Probability Calculators

I.A. Arrow-Debreu contracts

The most direct assessments of the probabilities of specific countries joining a European currency union are the prices of Arrow-Debreu securities that pay off contingent upon the countries being admitted. Two such contracts, on Italy and on Spain, began trading on the Iowa Electronic Markets on March 3, 1998. Each contract paid off $1 conditional upon Italy or Spain being selected by the European Council as eligible to participate in the currency union -- which transpired at the summit meeting on May 2. Market prices and volumes for the two contracts are graphed below, in Figure 2.

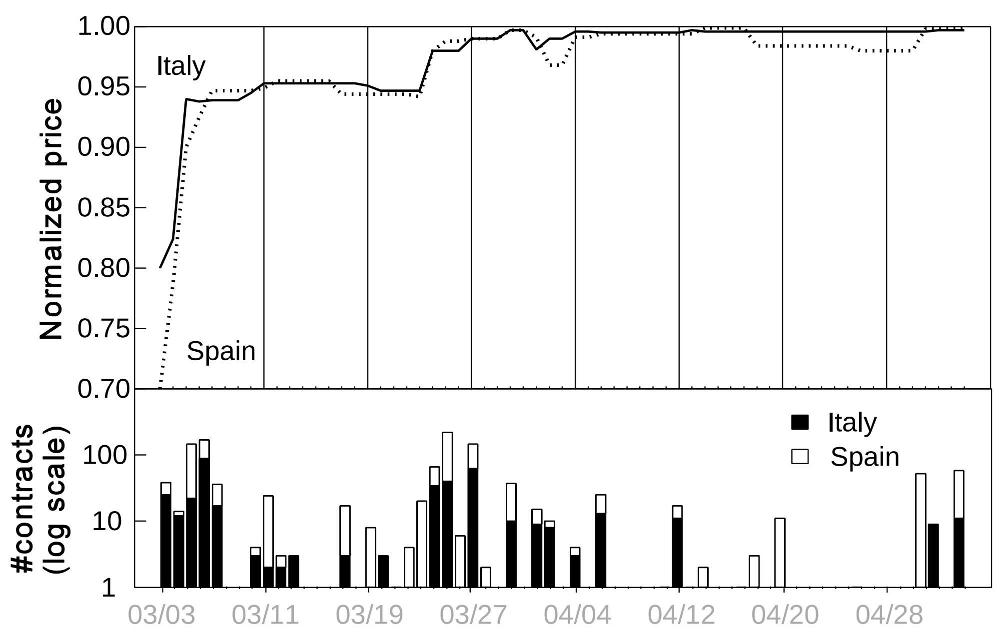

Figure 2. EMU Membership contracts: price and volume. Prices are for contracts that pay off $1 conditional upon admittance. Volume figures are the total number of country-specific contracts traded per day -- both “in” and “out” contracts.

Unfortunately, the contracts were introduced too late in the EMU process to contain much insight into the alternate methods of assessing EMU probabilities. The contracts actually began trading shortly *after* the European governments released their official economic results for 1997. All countries except Greece that were interested in participating in the currency union met the interest rate, inflation and budget deficit criteria of the Maastricht treaty, while exemptions on the public debt criterion for heavily indebted Italy and Belgium were widely expected on the grounds of substantial improvement.[4] Consequently, the prices on Italian and Spanish admission rapidly gravitated to 95 cents on the dollar once trading began in earnest on the contracts. Further substantial trading volume and price movements were observed prior to the March 27 official reports of the European Monetary Institute and the European Commission regarding countries' eligibility, at which time prices moved to 99 cents on the dollar

I.B Options-based EMU assessments

An alternate direct method of assessing EMU probabilities is to exploit the information from currency options' implicit volatilities regarding future intra-European cross exchange rate volatility. Stochastic volatility option pricing models such as Scott (1987) and Hull and White (1987) indicate that the "implicit" variance that equates the Garman-Kohlhagen (1983) currency option pricing formula to observed option prices should roughly be the expected average variance of exchange rate percentage changes over the lifetime of the option:[5]

$$\hat{\sigma}^2_{t,T} \approx \frac{1}{T} E_t \int_t^{t+T} Var_\tau\left(\frac{dS}{S}\right) d\tau \qquad \textbf{(1)}$$

[4]*New York Times*, "Europeans Clear Remaining Hurdle to Currency Unity." February 28, 1988,p.1ff.

[5]Several issues are being ignored in this simplified exposition: a concavity bias that (slightly) biases downward at-the-money implicit variances relative to expected average variance, the impact of volatility risk premia, and Hull and White's assumption that asset return shocks and volatility shocks are uncorrelated. For further discussion of these issues, see Bates (1996, pp. 590-91).

Furthermore, currency options' implicit volatilities do appear to forecast future exchange rate volatility reasonably well in general -- in contrast to those from stock or stock index options.[6] Consequently, setting implicit *forward* variances

$$\hat{\sigma}^2_{t,T_1 \to T_2} \equiv \frac{1}{T_2 - T_1} E_t \int_{t+T_1}^{t+T_2} Var_\tau \left(\frac{dS}{S} \right) d\tau = \frac{\hat{\sigma}^2_{t,T_2} T_2 - \hat{\sigma}^2_{t,T_1} T_1}{T_2 - T_1} . \qquad \textbf{(2)}$$

equal to the probability-weighted average of the zero value expected conditional upon currency union, and a higher estimated value conditional upon EMU not occurring, would appear a natural and direct way of assessing the probability of EMU transpiring.

The major difficulty with this approach is data availability. Interbank options on pound/DM, DM/FF, and DM/lira are actively traded, but primarily only for maturities of up to one year.[7] International convertible bonds such as those studied by Jennergren and Näslund (1990) often contain longer-maturity intra-European cross-rate options, but extracting implicit option prices is severely complicated by the plethora of bond-specific features. Consequently, the effective absence until 1998 of data on actively traded options maturing *after* January 1, 1999 has precluded the construction of EMU calculators based upon intra-European currency options. Some studies have, however, used such options to assess developments in the run-up to the currency union; notably Campa, Chang, and Reider (1997), Butler and Cooper (1997), and Adão, Cassola, and Luíz (1998).

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[6]Most empirical work on currency options has used the $/FC options traded on centralized exchanges. Those studies typically find implicit volatilities are roughly unbiased forecasts of future volatility; see Bates (1996) for a survey. The over-the-counter intra-European cross-rate options offered by many banks have been less actively studied, because of difficulties in acquiring data.

[7]The Bank for International Settlements (1996, Table D-8) reports that turnover in pound/DM over-the-counter options averaged about $1½ billion/day in April 1995, while FF/pound options' turnover was about $2 billion/day. However, Table D-6 indicates only 11% of all OTC currency options had maturities greater than 1 year.

Even if the data were available, implicit volatility-based EMU calculators would suffer from many of the difficulties of the interest rate-based calculators discussed below. The major issue is estimating what cross-rate volatility would be conditional upon a specific country failing to join the EMU. If, for instance, non-members nevertheless fix their exchange rates against the Euro-bloc countries (Butler and Cooper's ERM-2 scenario), the EMU and non-EMU cross-rate volatilities are virtually identical and EMU probability computations are infeasible. Nevertheless, using currency options to assess shifts in exchange rate regimes is a more direct approach than inferring shifts in regimes from bond yields, and it is regrettable that the data are not readily available.

II. Term structure-based EMU probability calculators

II.A. Theoretical foundations

The equilibrium approach to bond pricing exemplified by Cox, Ingersoll, and Ross (1985a,b) identifies zero-coupon bond prices of maturity T as

$$\begin{aligned} B(t,\ T) &= E_t[\tilde{M} \exp(-\bar{r}_{t,T} T)] \\ &\equiv E_t^*[\exp(-\bar{r}_{t,T} T)]\ , \end{aligned} \qquad \textbf{(3)}$$

where M is a positive pricing kernel with mean 1, $\bar{r}_{t,T} \equiv \frac{1}{T}\int_t^{t+T} r_{t+s}\,ds$ is the (random) average interest rate over [t, t+T], and E_t^* is a "risk-neutral" probability measure defined by (3) that incorporates the appropriate compensation for interest rate risk. The continuously compounded yield to maturity $y_{t,T}$ and instantaneous forward rate $f_{t,T}$ of maturity T are defined implicitly by

$$B(t,\ T) \equiv \exp(-y_{t,T}\,T) = \exp\left[-\int_t^{t+T} f_{t,\tau}\,d\tau\right]. \qquad \textbf{(4)}$$

Equivalently, the forward rate $f_{t,T}$ contractible now for instantaneously depositing or borrowing T periods hence is given by

$$\begin{aligned} f_{t,T} &\equiv -\frac{\partial \ln B(t,\ T)}{\partial T} \\ &= \frac{E_t^*\left[\tilde{r}_{t+T}\exp(-\bar{r}_{t,T}T)\right]}{E_t^*\left[\exp(-\bar{r}_{t,T}T)\right]} \\ &= E_t^*[\tilde{N}\tilde{r}_{t+T}] \\ &\equiv E_t^{**}[\tilde{r}_{t+T}]\,, \end{aligned} \tag{5}$$

where E_t^{**} is an alternate probability measure constructed analogously to the transformation in (3), using the Radon-Nikodym derivative $\tilde{N} = e^{-\bar{r}T} / E_t^*[e^{-\bar{r}T}]$. (5) can alternately be derived from the equilibrium condition that the "risk-neutral" expected discounted profit from forward rate speculation be zero:

$$E_t^*\left[e^{-\bar{r}_{t,T}T}(\tilde{r}_{t+T} - f_{t,T})\right] = 0\,. \tag{6}$$

The basic EMU probability calculator approach implemented *inter alia* by De Grauwe (1996), J.P. Morgan (1997) and Favero, Giavazzi, Iacone, and Tabellini (1997) uses the interest rate forecasts inferred from intra-European forward rates to assess the probability of any two countries entering into a currency union. For instance, if both Germany and Italy enter into a permanent currency union, the elimination of currency risk implies that DM and lira Italian money market deposits and loans become perfect substitutes, with identical interest rates. By contrast, a failure to achieve currency union implies a gap, presumably positive, between Italian and German interest rates. The Marshallian expectations hypothesis $f_{t,T} = E_t(r_{t+T})$, combined with a particular estimate of the expected future interest rate spread conditional upon EMU *not* occurring (NEMU), yields a method of inferring the probability of EMU occurring from observed forward rate differentials:

$$f_{t,T}^{ITL} - f_{t,T}^{DEM} = \pi_{EMU} \times 0 + \pi_{NEMU} E_t[r_{t+T}^{ITL} - r_{t+T}^{DEM} \mid NEMU]\,. \tag{7}$$

The alternate EMU probability calculators diverge primarily in their method of estimating the no-EMU scenario for future interest differentials, and in the selection of the future time interval.

Figure 3 shows the forward spreads for 7 currencies and for the ECU currency basket. Instantaneous forward rates corresponding to fixed 1999, 2002, and 2005 future maturities were computed from 3-, 6-, and 12-month Eurocurrency rates and 1-10 year swap rates, using Svensson's extension of the Nelson-Siegel (1987) methodology as described in Söderlind and Svensson (1997).[8] Forward spreads over 1991-98 were predicting relatively narrow post-1998 interest rate spreads relative to Germany for Belgium, Denmark, the ECU, France, and the Netherlands, but substantially larger spreads for Italy, Sweden, and the U.K. The forward spreads for most of the countries participating in the currency union (Belgium, France, and the Netherlands) had converged to zero by or before the fixing of official bilateral rates against the DM at the May 1998 summit. Interestingly, however, both the Italian and the ECU 1999 forward spreads were still slightly positive as of the summer of 1998. The latter can be attributed to exchange rate uncertainty regarding the rate of conversion of the ECU into euros, given the presence in the currency basket of non-EMU currencies such as the pound. The positive Italian spreads suggests that some uncertainty also remains as to whether the official lira/DM bilateral rate will in fact be achieved -- an uncertainty reflected in an unsuccessful speculative attack on the lira in August.[9]

Of the three countries (Denmark, Sweden, and the U.K.) that chose not to participate in the EMU in 1999, two (Denmark and Sweden) nevertheless exhibited substantial convergence in post-1998 forward rates relative to Germany over 1997 and 1998. The convergence was even more pronounced for Swedish 2002 and 2005 forward spreads, suggesting the possibility of delayed entry.

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[8]Further information on the data and the estimation methodology are in the appendix.

[9]*New York Times,* "Euro: Prenatal Force to Contend With", November 4, 1998.

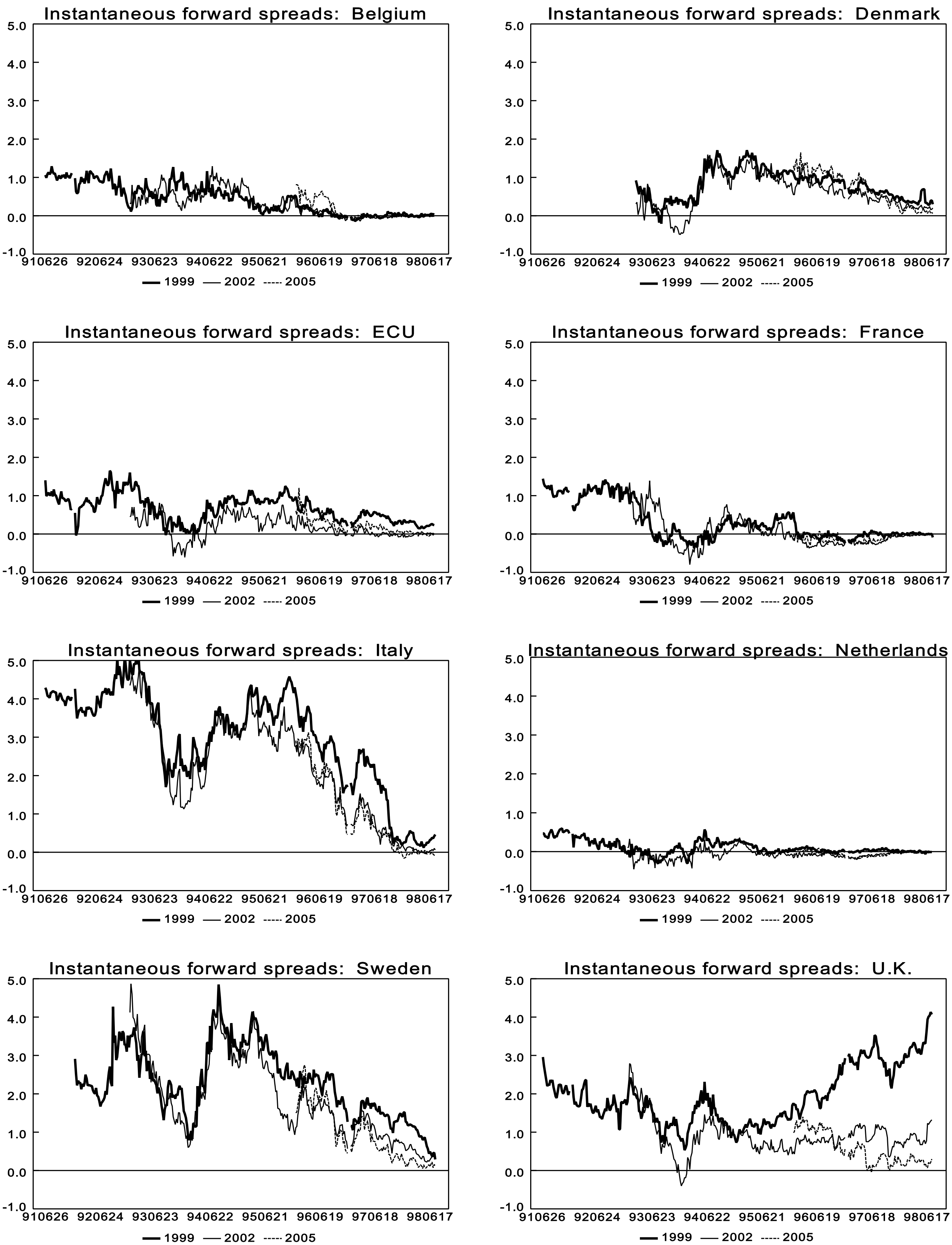

Figure 3. Instantaneous forward spreads relative to Germany.

British 1999 forward rates by contrast diverged increasingly over 1995-98 from German rates, but a sharply inverted forward spread curve over 1997-98 also suggests the possibility of delayed entry.

An immediate problem for EMU probability calculators is that although Germany is assumed to be the low-interest rate country, as has historically been the case within the European Union over 1977-98, forward spreads relative to Germany have occasionally been *negative* by 0-40 basis points for Denmark, France, the Netherlands, the U.K., and the ECU. This implies that despite history, market participants are assigning some probability to a post-1998 non-EMU scenario in which Germany is no longer the low-interest country. EMU probability calculators influenced by historical norms that calculate positive interest spreads conditional on EMU not occurring consequently estimate negative π_{NEMU} values over these periods, and π_{EMU} values greater than 1. De Grauwe's and Morgan's EMU calculators set such π_{EMU} values to 1, under the interpretation that they indicate a strong probability of EMU. Theoretically, however, a negative forward rate spread can constitute as strong evidence against future interest rate convergence as a positive spread.

Three issues arise in the assessment of whether EMU probability calculators are likely to do a good job. First is the issue of whether European term structures of interest rates have in fact demonstrated any ability to forecast future intra-European interest rate differentials. Second, there are assorted convexity biases, risk adjustments, and currency adjustments that can potentially bias inferred probabilities of monetary union away from the true conditional probabilities. Finally, there is the key issue of identifying the non-EMU regime.

II.B. The expectations hypothesis

EMU probability calculators are useless if European term structures contain little information regarding future interest rate changes. However, as discussed in Gerlach and Smets (1997a,b), Hardouvelis (1994), and Bekaert, Hodrick and Marshall (1997), European term structures tend to be more consistent with the expectations hypothesis than has been the U.S. experience. Table 1A reports country-specific estimates of the forward expectations hypothesis

$$R_{t+n}^{n} - R_{t}^{n} = \alpha + \beta[F(t, t+n \rightarrow t+2n) - R_{t}^{n}] + \varepsilon_{t+n} \qquad \textbf{(8)}$$

for n-month Eurodeposit rates R_t^n (n = 3, 6) relative to the forward rates F inferred from n-month and 2n-month Eurodeposit rates. Short-maturity forward rates are definitely informative with regard to future short-term rates, but the unbiasedness hypothesis is rejected for about half of the countries. Table 1B shows that intra-European forward *spreads* relative to the German forward rates do even better when forecasting of future interest rate differentials, with only 3 regressions out of 18 rejecting the unbiasedness hypothesis.

As discussed in Gerlach and Smets (1997a,b), part of this greater consistency with the expectations hypothesis is undoubtedly attributable to the greater past predictability of European interest rate changes. Speculative attacks on the weaker currencies participating in the European Monetary System generate temporarily high short-term interest rates for those currencies, and sharply inverted term structures that correctly predict future interest rate declines once the currency is devalued.[10] The expectations hypothesis fares less well for the DM, pound, and dollar, which were typically less constrained by exchange rate targets over 1977-98.

A past ability of term structures to forecast largely predictable changes in interest rates primarily associated with currency crises does not guarantee good forecasting performance for the future. A currency union is a European exchange rate regime previously experienced only by Belgium and Luxembourg, while the interest rate consequences of failing to achieve union could be

[10] Gerlach and Smets (1997b) note that while these crises are unquestionably important in tests of the expectations hypothesis, the hypotheis is also not rejected in more tranquil periods.

difficult to forecast. Nevertheless, it is reassuring that there are no strong grounds *a priori* for rejecting the use of forward spreads as a forecast of future interest rate differentials.

II.C. Divergences between inferred and conditional probabilities

As indicated above in equations (3) - (5), equilibrium bond pricing theory does not in general predict that forward rates should be unbiased predictors of future short-term interest rates. There are three potential sources of bias. First is the interest rate risk premium (or term premium) when going from conditional to "risk-neutral" bond pricing probability measures. Second is the use of forward rates rather than bond prices, inducing biases indicated in equation (4). Third is a currency conversion issue when comparing forward rates from different currencies.

According to equation (3), EMU probabilities inferred from bond prices are "risk-neutral" (or risk-adjusted) probabilities computed under the probability measure E_t^*:

$$B(t,\ T)\ =\ \pi_{EMU}\, E_t^*[e^{-\bar{r}T}\ |\ EMU]\ +\ \pi_{NEMU}\, E_t^*[e^{-\bar{r}T}\ |\ NEMU] \qquad \textbf{(9)}$$

Equivalently, $\pi_{EMU}\ \equiv\ Prob_t^*[EMU]$ is the current futures price on an Arrow-Debreu contract that pays off conditional upon EMU occurring. If the difference between EMU and no-EMU is associated with a major difference in investment opportunities, hedging against those differences could create a significant divergence between conditional probabilities and Arrow-Debreu futures prices.

Whether European Monetary Union will have real economic effects, and of what sign, has been a matter of considerable debate. Bean (1992), for instance, argued that both the long-term costs and the long-term benefits are likely to be small, implying little divergence between actual and risk-neutral probabilities. On the other hand, the short-term consequences for economic policies and interest rates of failing to achieve membership in a currency union could in principle be substantial.

For instance, a failure by Italy to achieve either the interest rate or inflation criteria of the Maastricht treaty might well have caused yet tighter monetary policy, in order to qualify at a later date. Whether this matters for risk premia does, however, depend upon the degree to which Italian bond markets are segmented from world markets. It appears appropriate to interpret inferred probabilities as "risk-neutral" probabilities and, to a first approximation, as conditional probabilities as well.

A crude calibration of the potential difference between actual and "risk-neutral" probabilities, inspired by Hansen and Jagannathan (1991) and Cochrane and Saa-Requejo (1996), can be obtained by placing bounds on the maximal feasible Sharpe ratio from speculating in an Arrow-Debreu forward contract on EMU occurring. If π is the futures price on EMU and p is the true probability, this approach implies $|p - \pi| \leq sr_{\max}\sqrt{p(1-p)} \leq \frac{1}{2} sr_{\max}$. Using $sr_{\max} = .3\sqrt{T}$ (the historical Sharpe ratio on the U.S. stock market) implies that actual and risk-neutral probabilities should deviate by less than 15% at an annual forecast horizon. However, this approach is too imprecise at longer horizons, indicating tighter theoretical bounds on $sr_{\max}$ are needed.

Forward rates versus bond prices

The second issue is the bias induced by using forward rates instead of bond prices when inferring (risk-neutral) probabilities of EMU occurring. Bond prices can be written as

$$\begin{aligned} B(t,\ T) &= E_t^*[e^{-\bar{r}T}] = \pi_{EMU} E[e^{-\bar{r}T} | EMU] + \pi_{NEMU} E[e^{-\bar{r}T} | NEMU] \\ &\equiv \pi_{EMU} B_{EMU} + \pi_{NEMU} B_{NEMU} . \end{aligned} \tag{10}$$

Consequently the forward rate is

$$
\begin{aligned}
f_{t,T} &\equiv -\frac{\partial \ln B(t,\ T)}{\partial T} \\
&= \frac{\pi_{EMU}\, E_t^{*}[r_{t+T}\, e^{-\bar{r}T} \mid EMU] \ +\ \pi_{NEMU}\, E_t^{*}[r_{t+T}\, e^{-\bar{r}T} \mid NEMU]}{B} \\
&= \pi_{EMU}\left(\frac{B_{EMU}}{B}\right) f_{EMU} \ +\ \pi_{NEMU}\left(\frac{B_{NEMU}}{B}\right) f_{NEMU} \\
&\equiv \pi^{*}_{EMU}\, f_{EMU} \ +\ \pi^{*}_{NEMU}\, f_{NEMU}
\end{aligned}
$$

where $f_{EMU} \equiv E_t^{*}[r_{t+T}\, e^{-\bar{r}\,T} \mid EMU] / B_{EMU}$ is the current forward rate *conditional* upon EMU occurring, and f_{NEMU} is defined similarly. The former is typically assumed equal to the current DM forward rate; the latter is estimated by various methods discussed in section II.C below. Divergences in bond prices conditional upon EMU occurring or not occurring will yield biased inferences of the probability of currency union from the current observed forward rate $f_{t,T}$.

How severe is this bias relative to risk-neutral probabilities inferred from bond prices? Re-expressing in terms of bond yields,

$$
\begin{aligned}
\pi^{*}_{EMU} &\equiv \pi \frac{B_{EMU}}{B} \\
&= \pi \frac{e^{-y_{EMU}T}}{\pi e^{-y_{EMU}T} + (1-\pi)e^{-y_{NEMU}T}} \\
&\approx \pi + \pi(1-\pi)(y_{NEMU} - y_{EMU})T\,.
\end{aligned}
\qquad \textbf{(12)}
$$

Thus, the bias depends upon the maturity of the forward rates used in assessing EMU prospects, and the divergence between the EMU and non-EMU contingencies. If, for instance, 1999 forward rates are used, then bond yields correspond to 1999 maturities and differ only insofar as the event of EMU occurring or not is associated with divergences in *pre*-1999 interest rate paths. For instance, Italy might be more likely to join the EMU under tighter pre-1999 monetary policies, implying $y_{EMU} > y_{NEMU}$ and a downward bias in inferred EMU probabilities.

If forward rates from later maturities are used, as in Favero *et al* (1997) and as graphed above in Figure (3), alternate biases arise. Bond yields for longer maturities can be decomposed into the 1999 bond yield and post-1998 average forward rates:

$$y_{t,T}\,T \;=\; y_{t,\,1999\,-\,t}(1999 - t) + \bar{f}_{1999\rightarrow T}(T - 1999)\;. \qquad \textbf{(13)}$$

Assuming for calibration purposes that the term structures conditional on EMU occurring or not occurring differ only in the average forward rates, EMU probabilities π^{*}_{EMU} inferred from forward rates are biased *upward* by the amount

$$\pi^{*}_{EMU} \approx \pi_{EMU} + \pi_{EMU}(1 - \pi_{EMU})\left[\bar{f}_{NEMU} - \bar{f}_{EMU}\right](T - 1999)\,. \qquad \textbf{(14)}$$

If the difference between EMU occurring and not occurring implies a difference of 4% in the average forward rates, using instantaneous forward rates for the year 2002 would bias inferred probabilities upward by at most $¼(.04)(3) = .03$. Thus, using forward rates instead of bond prices does not appear to significantly influence EMU probability computations.

Currency issues

A further potential bias originates in numeraire issues when considering the forward expectations hypotheses underlying EMU probability calculators. The “risk-neutral” expectational operators that incorporate the assorted risk premia by changing the probability measure do depend upon which currency is used when assessing investments. Consequently, forward rates from Italian and German swap rates cannot be compared without first expressing the underlying investments in a common currency.

The "risk-neutral" equilibrium condition that future currency speculation is currently expected to yield zero discounted profits indicates that

$$E^{*}\left[e^{-\bar{r}_{t,T}T}\left(r^{*}_{t+T}\,dt + \left(\frac{dS}{S}\right)_{t+T} - r_{t+T}\,dt\right)\right] = 0\,, \qquad \textbf{(15)}$$

where r^{*}_{t+T} is the future foreign interest rate and S is the exchange rate. Equivalently,

$$\frac{E_t^{**}(dS/S)_{t+T}}{dt} = E_t^{**}(r_{t+T}) - E_t^{**}(r^{*}_{t+T}) \qquad \textbf{(16)}$$

where E_t^{**} is the risk-adjusted expectations operator relevant to the *domestic* investor when assessing future speculations, given the domestic discount factor used in (15). From (5) above, $f_{t,T} = E_t^{**}(r_{t+T})$. However, the same is not true for the relationship between the foreign forward interest rate and future spot rate from the domestic investor's perspective. The zero-profit equilibrium condition on contracting to borrow foreign currency at the forward interest rate and subsequently investing at the spot interest rate is

$$E_t^{**}[S_{t+T}(\tilde{r}^{*}_{t+T} - f^{*}_{t,T})] = 0\,, \qquad \textbf{(17)}$$

which implies that the foreign forward rate is

$$f^{*}_{t,T} = E_t^{**}(r^{*}_{t+T}) + \frac{Cov_t^{**}(S_{t+T},\, r^{*}_{t+T})}{E_t^{**}(S_{t+T})}\,. \qquad \textbf{(18)}$$

Consequently, the relationship between forward spreads and future expected exchange rate changes when using the domestic currency as numeraire is

$$
\begin{aligned}
f_{t,T} - f^{*}_{t,T} &= \frac{E_t^{**}(dS/S)_{t+T}}{dt} - \frac{Cov_t^{**}(S_{t+T},\, r^{*}_{t+T})}{E_t^{**}(S_{t+T})} \\
&= \pi^{**}_{EMU} \times 0 + \pi^{**}_{NEMU} \frac{E_t^{**}\left[(dS/S)_{t+T} \mid NEMU\right]}{dt} - \frac{Cov_t^{**}(S_{t+T},\, r^{*}_{t+T})}{E_t^{**}(S_{t+T})} \, .
\end{aligned} \tag{19}
$$

I know of no EMU probability calculator that considers this potential source of bias. Papers that test or use the intra-European uncovered interest parity hypothesis at maturities of less than a year typically rule out a substantial effect from numeraire issues from the fact that intra-European exchange rate volatilities are small.[11] However, the maturities in (17) and (18) are substantially longer, creating the possibility of a greater impact.

II.C. Estimating the non-EMU scenario

One of the most critical issues in identifying the probability of two countries joining in a currency union is estimation of the relevant forecast conditional upon EMU *not* occurring. As this scenario represents a counterfactual hypothesis, the merits of the forecasts cannot directly be tested, but can only be judged by the overall merits of the methodologies. Three quite different approaches will be discussed here: J.P. Morgan's "kitchen sink" regression, Favero *et al*'s central bank reaction function, and Lund's term structure model. A fourth approach used by De Grauwe (1996) and Credito Italiano employs the average spreads from a period with low EMU prospects. De Grauwe uses 1990 average spreads, while Credito Italiano uses 1993.[12] Other banks have also created EMU calculators, but Morgan's is one of the few to publicly document its methodology.

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[11]See, e.g., Bertola and Svensson (1993).

[12]Favero *et al* (1997), p. 15.

J.P. Morgan's EMU calculator identifies non-EMU forward rate spreads by regressing daily 10-year swap spreads (the difference, e.g., between Italian and German swap rates) on a set of non-European interest rate variables intended to capture "international factors relating to investors appetite for risk or the supply of liquidity."[13] Non-European variables were used for exogeneity reasons, and the regression interval 1988-1992 was selected as a period in which expectations of future currency union were plausibly not affecting intra-European interest differentials. Regressions were only run through December 1991 for those countries that experienced currency devaluations in the fall of 1992. Various regressors are considered; the analysis concludes "[w]e are principally left with using the average of the U.S.-Canada and Japan-Australia 10yr bond spread and the U.S. and Japanese 10yr-2yr yield spread." The estimated "non-EMU" 10-year swap spread is then transformed into an estimate of the post-1998 forward swap spread by using the observed pre-1999 swap spread and the average slope of the swap spread curve over 1988-92.

While the objective of proxying for global trends in credit markets is laudable, the diagnostics of model performance in- and out-of-sample overstate the model's reliability. The major difficulty is that the independent and dependent variables, while stationary, are highly persistent. The high reported R^2's (ranging from 61% to 83% on 1300 daily observations over 1988-92) are therefore to be expected, while good out-of-sample forecasting performance over 1993 is also to be expected for highly persistent variables. Dickey-Fuller tests for stationarity of in-sample residuals are also run -- the standard Engle-Granger test for cointegrated variables -- but since the variables are stationary the relevance of this test is not apparent.

It does appear that the global variables used in assessing Morgan's non-EMU scenario contain some information for post-1992 swap spreads. In particular, several countries experienced temporary declines in swap spreads relative to Germany over 1992-94, and again after 1995 (see

---

[13]Morgan (1997), p. 4.

Figure 3) that are in part correlated with Morgan's estimated non-EMU swap spreads.[14] Overall, the regression approach appears somewhat preferable to just using average spreads over some interval as an estimate of the non-EMU scenario.[15] Nevertheless, it must be recognized that the non-EMU estimate potentially contains considerable structural instabilities that could affect EMU probability inferences.[16]

A further issue emphasized in Favero *et al* (1997), is that Morgan's use of the post-1998 section of 10-year swap spreads ignores the possibility of delayed EMU entry. For instance, a 10-year Italian-German swap spread as of January 1, 1995 incorporates instantaneous forward spread predictions of future interest differentials over 1999 through 2005. Morgan's EMU probability estimates should therefore be loosely interpreted as the probability of EMU entry on or some time after 1999.

Central bank reaction functions

Favero, Giavazzi, Iacone, and Tabellini (1997) use macroeconomic variables instead of financial variables in their forecast of the non-EMU scenario. Quarterly 3-month Euro-lira rates are regressed over 1987:I through 1996:II upon a lagged value, the inflation and growth gaps relative to Germany, a growth gap shift variable for the impact of German reunification, current and lagged log $/DM exchange rates, current and lagged 3-month Euro-DM interest rates, and dummy variable

---

[14]See Morgan (1997), Charts 3-9, which compare actual and non-EMU fitted swap spreads over 1988-97 for Italy, Belgium, Spain, Denmark, U.K., Sweden and France.

[15]Favero *et al* (1997) argue that Morgan's results for the lira are in fact very similar to those of Credito Italiano, which uses the average spread approach.

[16]Morgan (1997) argues that the results are insensitive to misspecification, and that a ±15% shift in the non-EMU forecast only shifts the inferred EMU probability roughly ±8% if the true probability is 50%. However, a 15% percentage error in a 4% swap spread forecast is only a 0.6% error absolute -- quite small compared to the forecast errors that might result from a misspecified regression equation and the observed large movements over 1988-98 in, e.g., Japanese bond yields.

that eliminates the impact of the 1992:IV interest rate outlier. The regression results are interpreted as representing the Bank of Italy's reaction function to macroeconomic fundamentals over 1987-96.[17] This regression, combined with current forecasts of the regressors,[18] is used to forecast future interest rates, and this forecast is used as the non-EMU scenario. The probabilities of Italy joining Germany in a currency union by 1999 or by 2001 are then inferred using (7) above and instantaneous forward rates computed from Euro-lira rates and lira swap rates via Svensson's extended Nelson-Siegel approach.

Favero *et al* emphasize that the decline in Italian forward rate spreads after 1996 was primarily attributable to improved fundamentals (e.g., falling Italian inflation) shifting the non-EMU forecast. The residual explanation of shifting EMU probabilities played less of a role. Their estimates also showed a relatively low probability over December 1995 through March 1997 of Italy joining the EMU in 1999, but a higher probability of it joining by 2002. The difference apparently reflects the sharply inverted term structure of forward rate spreads evident in Figure 3.

It seems likely that the reaction function estimated by Favero *et al* is partly capturing Italian monetary policy over 1987-1996. Achieving the Maastricht criteria certainly would create concern about inflation differentials, while the pre-1993 Exchange Rate Mechanism implies German interest rates were relevant over that period. However, whether forecasts based on this reaction function are relevant *conditional* upon Italy failing to enter EMU in 1999 is open to question. While the objective of qualifying by 2002 or later would plausibly prompt retention of the same reaction function, an alternate hypothesis is that the Bank of Italy might revert to its policies of the 1980's. The Bank of Italy reaction function estimated by Clarida, Galí, and Gertler (1997) over 1981:6

[17]Favero *et al* found no statistical evidence supporting interest rate sensitivity to a somewhat more traditional output gap measure: the deviation of actual GDP from a Hodrick-Prescott trend.

[18]The forecasts for inflation, output, and exchange rates for out to two years are based on "consensus" forecasts, while German interest rate forecasts are inferred from German forward rates.

through 1989:12 is quite different from that in Favero *et al*; in particular, higher steady-state real interest rates, and lower long-run sensitivity to Italian inflation, Italian output, and German interest rates. If this were used instead as the non-EMU scenario, inferred EMU probabilities would be quite different.

Bond pricing models

Lund (1998) provides a good example of the application of current bond pricing models to the issue of inferring EMU probabilities. Lund models, e.g., Italian instantaneous interest rates as the sum of the German rate and the Italian-German spread:

$$\begin{aligned} r_t^{ITL} &= r_t^{DM} + (r_t^{ITL} - r_t^{DM}) \\ &\equiv r_t^{DM} + s_t^{ITL} \end{aligned} \tag{20}$$

and makes the critical identifying assumption that the instantaneous spot rate spread evolves independently of the German interest rate. Under this assumption, plus the assumption that the exchange rate does not jump at EMU inception, he builds an explicit model of yield spreads. From (3) above, Italian bond prices are given by[19]

$$\begin{aligned} B^{ITL}(t,\ T) &= E_t^* \exp\left[-(\bar{r}_{t,T}^{DM} + \bar{s}_{t,T}^{ITL})T\right] \\ &= B^{DM}(t,\ T)\ E_t^* \exp\left[-\int_t^T s_u^{ITL}\, du\right] \end{aligned} \tag{21}$$

implying yield differentials between Italian and German discount bonds depend upon the expected future evolution of instantaneous spot rate differentials:

[19]Lund's derivation ignores the numeraire issues discussed in section II.B above. In particular, because of exchange rate risk prior to EMU, a lira-based risk-neutral expectation of the DM stochastic discount factor, $E_t^* \exp[-\bar{r}_{t,T}^{DM}]$, differs from the DM-based risk-neutral expectation, and consequently is not necessarily equal to the German bond price

$$
\begin{aligned}
S^{ITL}(t, T) &\equiv y_{t,T}^{ITL} - y_{t,T}^{DM} \\
&= -[\ln B^{ITL}(t, T) - \ln B^{DM}(t, T)]/T \\
&= \ln E_t^* \exp\left[-\int_t^T s_u^{ITL}\, du\right] .
\end{aligned}
\quad \textbf{(22)}
$$

If the two currencies join together in a currency union, the instantaneous spread drops instantly to zero. Rather than modeling this as only occurring on January 1, 1999, however, Lund posits a post-January 1999 *hazard rate* that allows for delayed entry. The current assessment θ_t of that hazard rate determines the current EMU probability assessment $[1 - \exp(-\theta_t \tau)]$ of a country joining Germany in a currency union within τ years after 1/1/99. Combined with assumed independent Ornstein-Uhlenbeck (or AR(1)) processes for the instantaneous spread s_t conditional upon no unification and the market price λ_t of spread risk, Lund develops a 3-factor model of yield spreads. In continuous time, the pre-EMU "risk-neutral" processes used in pricing spreads via (22) are

$$
\begin{cases}
ds = [\kappa_1(\mu_1 - s) + \kappa_1 \lambda]dt + \sigma_1 dW_1 - s\, dq \\
d\lambda = \kappa_2(\mu_2 - \lambda)dt + \sigma_2 dW_2 \\
d\theta = \sigma_3 dW_3
\end{cases}
\quad \textbf{(23)}
$$

where the W's are independent Wiener processes and q is a 0- or 1-valued indicator for the commencement of the currency union:

$$
Prob[dq_t = 1] = \begin{cases} 0 \;\; if \;\; t \le 1/1/99 \\ \theta_t dt \;\; otherwise . \end{cases} \quad \textbf{(24)}
$$

But whereas the yield spreads $S^{ITL}(s_t, \lambda_t, \theta_t, t, T)$ for different maturities are priced *as if* the state variables follow (23), the *actual* instantaneous spreads s_t (roughly equal to short-maturity Eurorate differentials) are assumed to follow the AR(1) process

$$
ds = \kappa_1(\mu_1 - s)dt + \sigma_1 dW_1 \quad \textbf{(25)}
$$

prior to January 1, 1999.

Lund estimates the parameters in (12) and the state variable realizations via nonlinear Kalman filtration, using two data sources: weekly 1-, 2-, 3-, 6-, and 12-month Eurorate differentials, and 1-10 year yield differentials inferred from swap rates. The broadest data interval was January 1990 through August 12, 1998, although subsets of that interval were used for specific countries either because of data unavailability, or because of deliberate data exclusion. First, short-maturity interest rates were excluded for currencies affected by speculative attacks prior to September 1993, so that the impact of a *fourth* underlying state variable (currency crises) could plausibly be ignored. Second, post-1999 maturities were not considered prior to 1995, because of lack of confidence in the inferred EMU hazard rates. The vector of yield spreads for different maturities provided cross-sectional evidence each period regarding parameter values and state variable realizations, while the time series evolution of those yield differentials provided further evidence regarding parameter values.

Lund makes two key identifying assumptions. First, he explicitly rules out the "double-decay" process of Jegadeesh and Pennacchi (1996) on the grounds of parsimony, and assumes that the "risk premium" state variable λ_t does not affect the *actual* AR(1) process (25) followed by the instantaneous spread. This assumption appears empirically reasonable for the 1990-98 period considered, as will be discussed below. However, one advantage of the more general model, as recognized in Andersen and Lund (1996), is that the second state variable can capture structural shifts in the nominal interest spread processes resulting, e.g., from shifts in inflation targets.

Second, the model implies yield spreads from bonds maturing *prior* to January 1, 1999 are unaffected by shifts in the EMU probability state variable θ_t. This is critical, because the "market price of risk" λ_t is also a free state variable that captures some movements in the shape of the term structure unjustified by the AR(1) process (25) estimated for the instantaneous spread s_t.[20][21] However, λ_t affects all maturities, whereas θ_t affects only post-1999 maturities. θ_t is inferred in essence by the difference between the observed post-1999 portion of the yield curve, and the post-1999 "local yield curve" predicted by extrapolating the pre-1999 yield curve for estimated state variables (s_t, λ_t). Thus, Lund's EMU probability assessments are essentially assuming that the non-EMU scenario for interest rate differentials is largely identified by the estimated 2-factor behavior of the term structure of yield spreads over 1990-98 for pre-1999 maturities.

Judging from Figure 1, Lund's model appears to estimate generally lower Italian-German interest differentials conditional upon EMU not occurring, and consequently higher non-EMU probabilities (lower EMU probabilities) than other models. Despite the low standard errors of Lund's structural parameter estimates, however, it is not really possible to say whether Lund's inferred non-EMU scenarios are better or worse than the Morgan and Favero *et al* approaches. The problem is that structural parameters are inferred partly from the time series behavior of the term structure of yield spreads, and partly from what is needed to match the cross-sectional patterns of yield spreads on individual weeks. The appropriate statistical theory for assessing the latter informational source has not really been developed. The typical null hypothesis also used here is that market participants know the true, time-invariant deep structural parameters and the state variable realizations with certainty when pricing bonds, and that observed and actual bond prices deviate only by independent, homoskedastic measurement error. Under this strong null hypothesis and given abundant cross-sectional data, inferred parameters tend to have low estimated standard errors.

[20]The 1-factor Vasicek model given by the risk-neutral equivalent of (14) primarily allows upward or downward sloped term structures that converge exponentially at longer maturities towards a fixed infinite-maturity yield; see Hull (1993, ch. 11). 2-factor models allow more complicated hump and U shapes, as well as more (state-dependent) variation over time in the slope and curvature of the term structure.

[21]The current bond pricing literature does not typically attempt to impose constraints on such risk premia from asset pricing models. See, e.g., Duffie and Singleton (1997).

Lund's estimates are best viewed as a model-specific description of a non-EMU scenario consistent with term structures of yield spreads observed over 1990-98. Term structures observed over the 1980's might yield different estimates of the non-EMU scenario.

Some insight into the time series stability of the term structure of spreads at short maturities can be gleaned from Eurocurrency spreads, for which data over a longer time interval are more readily available. Lund's model implies that the instantaneous spread is Markov, so the *slope* of the term structure of spreads contributes no additional information. A crude proxy for the former that partially avoids the impact on short maturities of speculative attacks is the six-month Eurocurrency spread; the latter is proxied by the difference between 12- and 6-month spreads. Table 2 reports estimates of the regression

$$s_t^6 = a + b\, s_{t-1}^6 + c\,[s_{t-1}^{12} - s_{t-1}^6] + \varepsilon_t \qquad \textbf{(26)}$$

for weekly data over 1977-1998, estimates for c over the 1990-98 subsample used by Lund, and heteroskedasticity-consistent Chow tests of process stability.

With the exception of the Dutch-German spread slope, all subsample estimates of $\hat{c}_{1990-98}$ are statistically insignificant, justifying Lund's specification. Over the longer 1977-98 interval, spread slopes are statistically significant for Denmark, Italy, and the Netherlands, but not for the other five countries. However, it does appear that including the spread slope captures parameter instabilities in the estimated spread process to some extent. *P*-values are substantially lower for subsample stability tests of Lund's AR(1) than for stability tests of (26), for 7 out of 8 countries. Parameter stability of (26) is rejected at the 10% level for 3 countries (Belgium, Great Britain, and Sweden), whereas the subsample stability of Lund's AR(1) is also rejected for these three, Denmark and Italy.

III. Prospects for the Future

There has been considerable speculation regarding what monetary policies will be pursued by the new European Central Bank (ECB). Some have argued that the ECB is likely to initially pursue a tight monetary policy, in order to establish an anti-inflationary reputation. The selection of Duisenberg as its initial head is certainly consistent with this perspective. Others have argued that the voting power of countries such as Italy that historically have been less concerned about inflation may result in monetary policy somewhat weaker than that historically pursued by the Bundesbank.

Morgan (1997) argues that the EMU probability calculators indicate a weak-ECB scenario. Fluctuations in the average EMU probability of three "periphery" countries (Italy, Spain, and Sweden) have been strongly correlated with deutsche mark weakness against the dollar. The argument is illustrated in Figure 4, which shows that the average Italian-Swedish 1999 forward spread against the DM has indeed fluctuated roughly in tandem with the $/DM exchange rate.

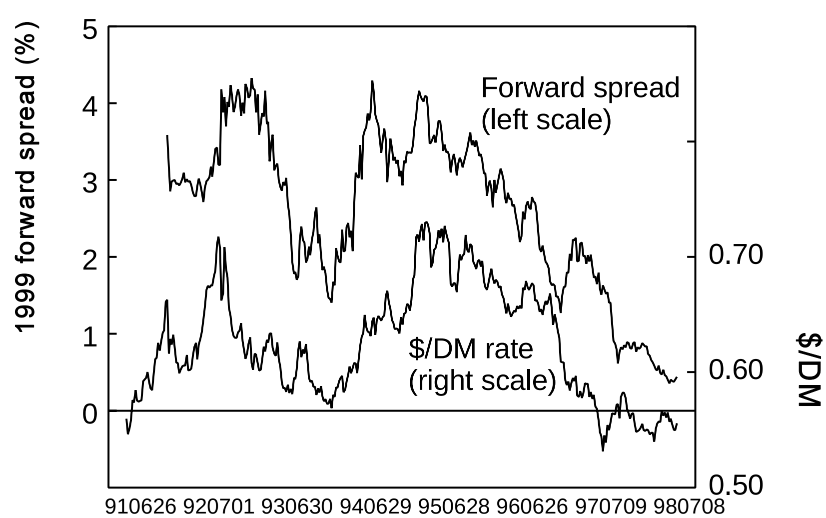

Figure 4. Average of Italian and Swedish 1999 forward spreads relative to Germany, and the $/DM exchange rate.

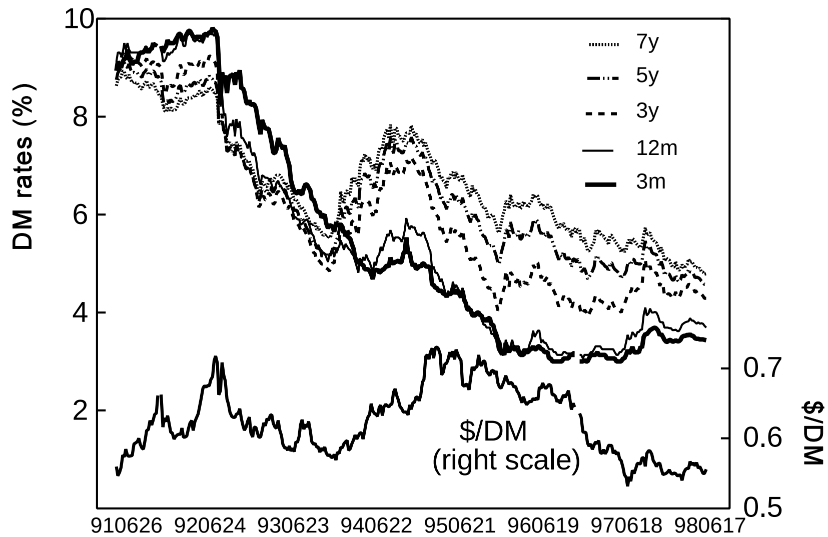

Figure 5. 3- and 12-month DM Eurodeposit rates, 3-, 5-, and 7-year forward rates, and the $/DM exchange rate.

Narrowing spreads (which EMU probability calculators would interpret as higher unification probabilities) have roughly coincided with a weakening DM, while widening spreads have coincided with a strengthening DM.

One difficulty with this argument is that forward spreads over this period are a good proxy for the overall level of European bond yields and forward rates, and for German rates in particular. Figure 5 shows German 3-month Eurodeposit rates and forward rates at various maturities over the same interval. Falling German rates were accompanied by even greater declines in non-German rates, while rising German rates accompanied larger increases elsewhere. The synchroneity was most pronounced for forward rates of at least 1 year in maturities. Consequently, in terms of the implications for the $/DM exchange rate, it is impossible to distinguish the weak-ECB hypothesis from an alternate hypothesis that the exchange rate fluctuations reflected German forward interest rate fluctuations over 1992-96 resulting, e.g., from shifting near-term expectations of future Bundesbank monetary policy.

Filtration-based assessments

As indicated above in Figure 3, forward rates for all countries participating in the currency union have currently essentially converged to a common "Euro" term structure. Some insights into financial markets' assessment of the future interest rate policies of the ECB *relative* to earlier Bundesbank policies can potentially be gleaned by comparing this yield curve with historical German norms. To this end, a state space representation of the vector of Eurocurrency deposit rates and swap rates for German instruments maturing *prior* to January 1, 1999 was estimated via Kalman filtration on weekly data:

$$\begin{aligned} \boldsymbol{y}_t &= \boldsymbol{A} + \boldsymbol{H}\boldsymbol{z}_t + \boldsymbol{u}_t, \quad \boldsymbol{u}_t \sim N(\boldsymbol{0}, \boldsymbol{R}) \\ \boldsymbol{z}_t &= \boldsymbol{F}\boldsymbol{z}_{t-1} + \boldsymbol{v}_t, \qquad \boldsymbol{v}_t \sim N(\boldsymbol{0}, \boldsymbol{Q}) \end{aligned} \tag{27}$$

where $\boldsymbol{y}_t$ is a 9×1 vector of 3-, 6-, and 12-month Euro-DM deposit rates and 1-7 year swap rates (excluding 6-year rates), expressed as continuously compounded yields; and $\boldsymbol{z}_t$ is a 4×1 vector of underlying state variables.

The Euro-DM data were available from January 5, 1997, while most of the swap rate data began in June 24, 1991.[22] Data for post-1998 maturities were treated as missing data, with only 3- and 6-month Euro-DM rates available for inference on the final date of July 1, 1998. The estimated state space model summarizes the time-series based information contained in the level and shape of the German term structure with regard to future term structure evolution, as well as the current assessment of the state vector *conditional* on pre-1999 maturities. The model was used to address two questions:

1. How does the current term structure of German swap yields compare with the German pre-1999 norm represented by the final filtered value $\hat{y}_{T|T}$?

[22] Further details on data, estimation methodology, and identifying restrictions are provided in the appendix.

2. How does the current term structure of German forward rates compare with the future 3-month spot rates that would be predicted based upon pre-1999 norms?

Figure 6 shows the actual and fitted yield curves, as well as 95% confidence intervals around the fits. The current DM yield curve is flatter than that which would be predicted based up historical German swap rates over 1991-1998 and given current short-term interest rates. The divergence is statistically significant at longer maturities. However, it should be recognized that the fits for longer maturities are based upon term structure relationships estimated on increasingly narrower subsets of the 1991-97 swap data interval, given the deliberate exclusion of swaps with payments beyond January 1, 1999.

Figure 7 presents a different picture of future interest rate prospects. The German 3-month interest rate is stationary but highly persistent, with a weekly autocorrelation of .9969 over 1977-98. Based upon current and recent German term structures for pre-1999 maturities, the state space model

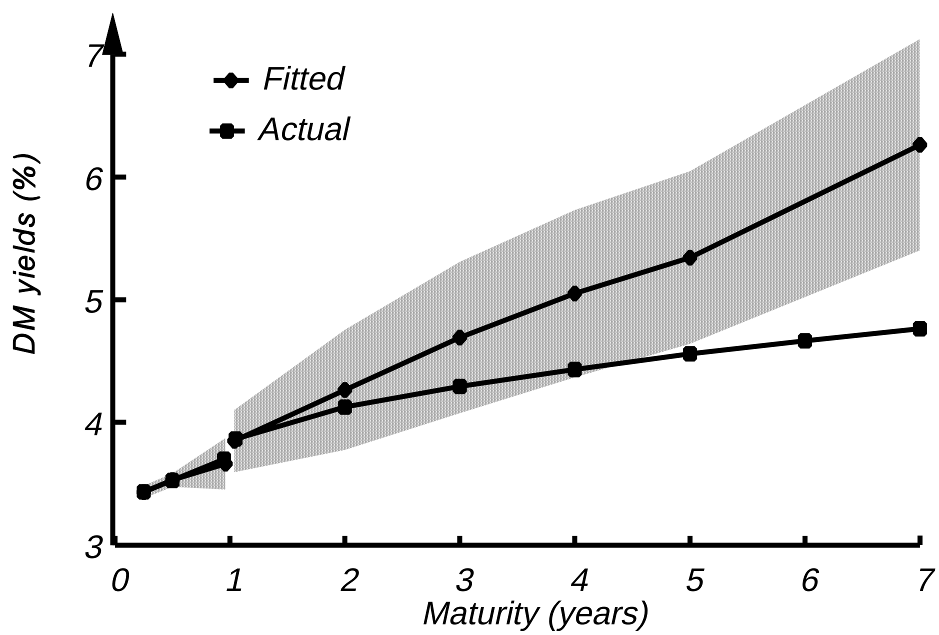

Figure 6. Actual and fitted continuously compounded DM yields on July 1, 1998. 3-12 month LIBID rates, and 1-7 year swap rates.

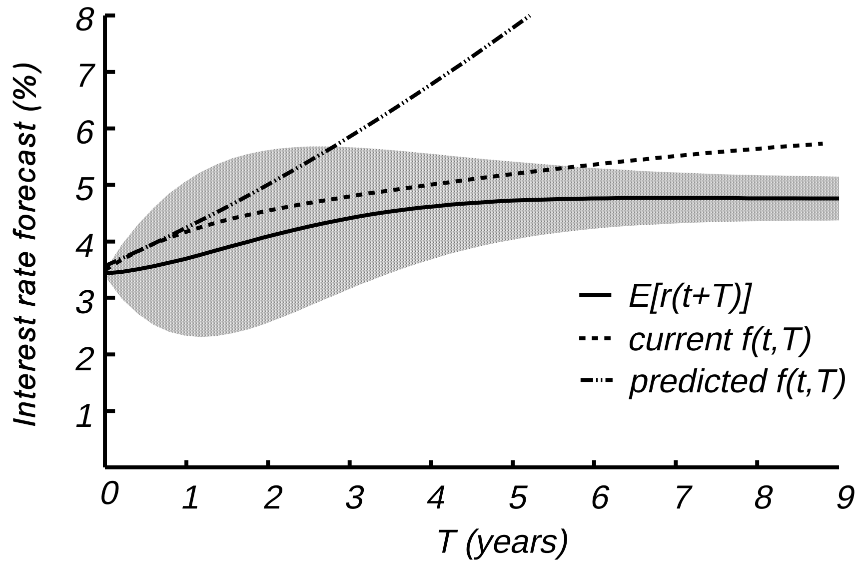

Figure 7. Alternate forecasts of future Euro interest rates. Forecasts are based upon the German state space model, on the current forward curve, and on the German forward curve predicted by the German state space model. Confidence intervals (shaded region) reflect uncertainty about the appropriate forecast.

would forecast gradual mean reversion in German rates back to a steady-state level of 4.75%. The current "unusually" flat Euro forward rate curve is largely consistent with that forecast. By contrast, the forward rate curve consistent with the fitted "German" yield curve of Figure 6 implicitly forecasts considerably faster increases in future German rates.

In summary, the current Euro term structure is unusually flat by German standards, but that flatness is consistent with how German short-term rates have historically evolved. This suggests that German term structures have historically been inconsistent with German short-term rate evolution (a hypothesis supported by the short-maturity rejections of German forward rate unbiasedness in Table IA above), but doesn't answer the question of how ECB monetary policies are likely to compare with past Bundesbank policies. At present, the only evidence we have is from exchange markets, and even that has multiple interpretations. The overall declines in German and European interest rates over 1996-98 have accompanied substantial weakening of the EMU currencies against the dollar. However, those could be interpreted either as a vote of no confidence in future ECB policies, or as an expectation of future economic recession.

Appendix

Data

Interest rate data were obtained from two sources. The Bank for International Settlements provided daily 3-, 6-, and 12-month Eurocurrency deposit rates for 21 countries, collected around 10 AM Swiss time. The data begin on January 3, 1977 for five currencies (German mark, Dutch guilder, Swiss franc, pound, and U.S. dollar), on September 1, 1977 for most other currencies, and run through August 31, 1998.

Swap rate data were downloaded from Datastream for nine European countries. Swap rates are the European coupon rates in a semiannual exchange of fixed European interest payments against floating dollar interest rate payments indexed to the Eurodollar interest rate; principal is also exchanged at maturity. Since the dollar floating-rate cash flows are at par at the swap's inception, swap rates provide the European coupon rate such that a European bond with semiannual coupons would be at par -- or, equivalently, the European semiannual yields to maturity, times two. Swap rate data are more readily available than country-specific bond yields. As with other Euromarket data, swap rates also alleviate concerns about country-specific default risk, capital controls, and tax issues, as well as eliminating heterogeneity across national data sources.

Midpoint swap rates were used; the average of bid and ask. Swap rates for 2-5, 7, and 10-year maturities were available for all countries except Denmark and Sweden beginning on June 24, 1991. Danish swap rates began on February 25, 1993, while Swedish rates began on January 2, 1992. Data for 1-, 6-, 8-, and 9-year maturities began on November 15 or 16, 1994 for all countries except France. French franc 1-year swap rates began on October 31, 1994, while 6-, 8-, and 9-year maturities began on January 24, 1995. Weekly (Wednesday) observations were selected from the daily data, with the nearest other day used when Wednesday was a holiday.

Forward rates

The 3- and 6-month forward rates used in testing the expectations hypothesis in section II.B were computed from the Eurodeposit rates using the formula[1]

$$1 + F[t,\ T_1 \to T_2]\ \frac{N_2 - N_1}{360} \equiv \frac{1 + R_{t,N2}(N_2/360)}{1 + R_{t,N1}(N_1/360)}\ . \qquad \textbf{(A.1)}$$

where N_1 (N_2) is the number of days until the shorter (longer) Eurodeposit matures. For 3-month forward rates, N_1 was roughly 91 days and N_2 was roughly 182 days, although there was some variation across months and because of bank holidays (identified by missing data) and weekends. This 3-month forward rate was used to forecast the future 3-month rate on a Eurodeposit that begins when the deposit corresponding to $R_{t,T1}$ ends -- i.e., the rate 2 business days before the first contract matures. Because of weekends and holidays, this second 3-month deposit did not necessarily mature precisely on the same day as the 6-month contract. This minor deviation in maturity was not considered a serious concern. 6-month forward rates were constructed analogously from 12- and 6-month Eurodeposit rates.

The instantaneous forward rates computed in section II.A used Svensson's extension of the Nelson-Siegel method, as described in Söderlind and Svensson (1997):

$$f_{t,T} = \beta_0 + \beta_1 \exp\left(-\frac{T}{\tau_1}\right) + \beta_2 \frac{T}{\tau_1} \exp\left(-\frac{T}{\tau_1}\right) + \beta_3 \frac{T}{\tau_2} \exp\left(-\frac{T}{\tau_2}\right) \qquad \textbf{(A.2)}$$

where β_0, $\beta_0 + \beta_1$, τ_1 and τ_2 are free positive parameters and β_2 and β_3 are unconstrained parameters. The continuously compounded yields for zero coupon bonds can be evaluated using the relationship

$$y_{t,T}\,T = \int_t^{t+T} f_{t,s}\,ds; \qquad \textbf{(A.3)}$$

analytic solutions are in Söderlind and Svensson (1997). The parameter vector for each day was estimated using 3 Eurodeposit rates and up to 10 swap rates, using the following loss function:

[1] See, e.g., Grabbe (1991, pp.264-5).

$$\ln L(\theta) \;=\; \min \sum_{n=1}^{3} \{\, y_{t,T_n} \;-\; [\hat{y}(T_n,\ \theta) \;-\; k] \}^2 \;+\; \sum_{n=4}^{N} \left[\frac{\ln \hat{B}(T_n,\ c_n;\ \theta)}{\tfrac{1}{2} D(T_n,\ c_n)} \right]^2 \qquad \textbf{(A.4)}$$

where

$y_{t,T_n},\ n = 1, \ldots 3$, is the continuously compounded yield computed from Eurodeposit rates;

$\hat{y}(T_n,\ \theta)$ is the fitted yield given parameter vector θ;

k is a free parameter used to splice together Eurodeposit and swap rate data;[2]

$\hat{B}(T_n,\ c_n;\ \theta)$ is the estimated price of a bond paying semiannual coupons at swap rate c_n over maturity T_n, given parameter vector θ; and

$D(T_n,\ c_n) \;=\; \dfrac{1 \;-\; (1 + \tfrac{1}{2} c_n)^{-2T_n}}{1 \;-\; (1 + \tfrac{1}{2} c_n)^{-1}}$ is the semiannual duration of a par bond with coupon rate c_n.

The loss function is essentially in yield space rather than in bond price space, reflecting the way the data are quoted. In particular, weighting log bond prices by the inverse of duration approximately transforms the last terms in (A.4) into the deviation between the observed swap rate c_n and the U.S.-style yield to maturity $\hat{Y}$ inferred from fitted bond prices:

$$\begin{aligned} \ln \hat{B}(T_n,\ c_n;\ \theta) \;&\approx\; \ln 1 \;+\; \left. \frac{\partial \ln \hat{B}}{\partial \ln(1 + \tfrac{1}{2} Y)} \right|_{Y = c_n} [\, \ln(1 + \tfrac{1}{2} \hat{Y}) \;-\; \ln(1 + \tfrac{1}{2} c_n) \,] \\ &\approx\; \tfrac{1}{2} D(T_n,\ c_n)(\, \hat{Y} \;-\; c_n \,) \;, \end{aligned} \qquad \textbf{(A.5)}$$

exploiting the fact that swap rates are coupon rates for bonds that are at par. All parameters were estimated using the Davidon-Fletcher-Powell nonlinear optimization algorithm (GQOPT subroutine DFP), with various parameter transformations used to enforce nonnegativity constraints.

---

[2]Figure 6 illustrates the significant gap between 1-year Eurodeposit rates and 1-year midpoint swap rates that necessitated the splicing parameter k.

Kalman filtration

A state space representation is of the form

$$\begin{aligned} \boldsymbol{y}_t &= \boldsymbol{A} + \boldsymbol{H}\boldsymbol{z}_t + \boldsymbol{u}_t, \quad \boldsymbol{u}_t \sim N(\mathbf{0}, \boldsymbol{R}) \\ \boldsymbol{z}_t &= \boldsymbol{F}\boldsymbol{z}_{t-1} + \boldsymbol{v}_t, \qquad \boldsymbol{v}_t \sim N(\mathbf{0}, \boldsymbol{Q}) \end{aligned} \qquad \textbf{(A.6)}$$

where $\boldsymbol{y}_t$, $\boldsymbol{A}$, and $\boldsymbol{u}_t$ are $N \times 1$, $\boldsymbol{z}_t$ and $\boldsymbol{v}_t$ are $K \times 1$, $\boldsymbol{R}$ is an $N \times N$ matrix assumed diagonal and nonnegative, $\boldsymbol{Q}$ is a $K \times K$ positive semidefinite matrix, and $N > K$. (A.6) is underidentified, since any transformation of the underlying state vector(e.g., $\boldsymbol{x}_t = \boldsymbol{M}\boldsymbol{z}_t$) constitutes an equally valid representation. Provided the $N \times K$ matrix $\boldsymbol{H}$ is of full rank K, the state space representation can be uniquely identified by imposing semi-arbitrary restrictions on any K rows of the matrix; for instance, setting $\boldsymbol{H}' = (\boldsymbol{I}_K \; \boldsymbol{H}_2')$, where $\boldsymbol{H}_2$ is an unconstrained $(N-K) \times K$ matrix. The system (A.6) can be estimated by Kalman filtration methods discussed in Shumway and Stoffer (1982), Watson and Engle (1983), and Hamilton (1994).

The data vector y_t consists of 3-, 6-, and 12-month DM Eurodeposit rates, and 1-7 year swap rates. All instruments maturing after January 1, 1999 were excluded, resulting in total elimination of the 6-year swap rates that only became available in 1994. Eurodeposit rates were available over an extended 1977-1998 interval; swap rates only after June 24, 1991. However, Kalman filtration methods can cope relatively easily with mismatched data sets and missing data.

The system (A.6) was estimated using various choices of the number K of factors. The first three identifying restrictions were chosen successively to mimic the level, slope, and curvature term structure factors advocated by Litterman and Sheinkmann (1988) -- albeit for the short-maturity *Eurodeposit* rates for which there was the most data. The fourth factor was chosen to mimic the slope of the term structure for longer maturities. As discussed above, these restrictions are arbitrary. Many of the factors exhibited near-unit root behavior. For instance, the single-factor model of the term structure had a weekly autocorrelation of .99772, reflecting high persistence in German interest rates. Consequently, a *conditional* Kalman filtration was estimated that included estimation of the initial state vector $\boldsymbol{z}_0$ as of December 27, 1976.

Table A.1 below summarizes the performance of the various models. While adding more factors continued to be highly significantly statistically up through five factors, out-of-sample forecasts for the five-factor model were more unstable. Furthermore, there were difficulties in inverting the information matrix when computing standard errors. Consequently, the 4-factor model was used.

Table A.1: Summary statistics for multifactor models.

# factors	# parameters	log likelihood	Eigenvalues of $\boldsymbol{F}$ (modulus)
1	29	4,373.35	.998
2	49	8,020.72	.997, .978
3	54	8,856.16	.997, .969, .969
4	68	9,033.24	.989, .989, .971, .743
5	83	9,117.60	.992, .992,. .931, .877, .736

The final forecasts $\hat{\boldsymbol{y}}_{T+n|T} = \boldsymbol{A} + \boldsymbol{H}\left(\boldsymbol{F}^n \hat{\boldsymbol{z}}_{T|T}\right)$ for July 1, 1998 and beyond were computed based upon the estimated process (A.6) and conditional upon pre-1999 maturity data, up through and including 3- and 6-month Eurodeposit rates observed on July 1. Standard errors were computed using the covariance matrix

$$Var\left(\hat{\boldsymbol{y}}_{T+n|T}\right) = \left.\frac{\partial \hat{\boldsymbol{y}}_{T+n|T}}{\partial \theta'}\right|_{\theta = \hat{\theta}} Var(\hat{\theta}) \left.\frac{\partial \hat{\boldsymbol{y}}_{T+n|T}}{\partial \theta}\right|_{\theta = \hat{\theta}} + \hat{\boldsymbol{H}}\boldsymbol{F}^n \boldsymbol{P}_{T|T} \left(\boldsymbol{F}^n\right)' \hat{\boldsymbol{H}}' \qquad \textbf{(A.7)}$$

where $Var(\hat{\theta})$ was estimated from the information matrix computed during optimization, and $\boldsymbol{P}_{T|T}$ was the estimated conditional covariance matrix of the final state variable realization. The first term captures the impact of parameter uncertainty; the second, the impact of uncertainty regarding the current state variable. It should be emphasized that the standard errors computed in this fashion reflect the uncertainty about the current *forecast* of future data; not the uncertainty about the future data. The purpose is to compare this forecast with other forecasts constructed using post-'99 maturities.

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Table IA

Tests of forward interest rate unbiasedness: $R_{t+n}^{n} - R_{t}^{n} = \alpha + \beta[F(t,\ t+n \rightarrow t+2n) - R_{t}^{n}]$

Country	Starting Date	#obs. (3-mth)	*n* = 3 months				*n* = 6 months				*P*-value: $\alpha=0, \beta=1$	
			α	*(s.e.)*	β	*(s.e.)*	α	*(s.e.)*	β	*(s.e.)*	3 mths	6 mths
Germany	770105	1117	-.02	*(.09)*	.53	*(.20)*	.05	*(.16)*	.60	*(.29)*	.054	.377
Belgium	770907	1082	.11	*(.12)*	.70	*(.17)*	.09	*(.19)*	.43	*(.25)*	.155	.058
Denmark	770907	1082	.20	*(.15)*	.79	*(.10)*	.17	*(.20)*	.59	*(.18)*	.047	.031
ECU	841107	708	.02	*(.07)*	.85	*(.14)*	.03	*(.16)*	.81	*(.25)*	.521	.962
France	770907	1082	-.03	*(.20)*	.97	*(.15)*	.21	*(.28)*	.77	*(.11)*	.964	.016
Italy	770907	1082	.18	*(.18)*	1.01	*(.14)*	.54	*(.28)*	.79	*(.10)*	.592	.012
Netherlands	770105	1117	-.02	*(.10)*	.84	*(.18)*	.04	*(.17)*	.85	*(.24)*	.683	.796
Sweden	770907	1082	.27	*(.12)*	.84	*(.12)*	.26	*(.20)*	.61	*(.19)*	.027	.035
U.K.	770105	1117	.10	*(.16)*	.59	*(.17)*	.28	*(.26)*	.61	*(.19)*	.004	.015
U.S.	770105	1117	-.03	*(.15)*	.76	*(.30)*	.06	*(.24)*	.43	*(.31)*	.687	.168

Table IB

Tests of forward *spread* unbiasedness (relative to German rates): $Rdif_{t+n}^{n} - Rdif_{t}^{n} = \alpha + \beta[Fdif_{t}^{n} - Rdif_{t}^{n}]$

Country	Starting Date	#obs. (3-mth)	*n* = 3 months				*n* = 6 months				*P*-value: $\alpha=0, \beta=1$	
			α	*(s.e.)*	β	*(s.e.)*	α	*(s.e.)*	β	*(s.e.)*	3 mths	6 mths
Belgium	770907	1082	.13	*(.11)*	.76	*(.19)*	.05	*(.11)*	.51	*(.29)*	.525	.231
Denmark	770907	1082	.19	*(.21)*	.76	*(.14)*	.15	*(.21)*	.69	*(.20)*	.027	.197
ECU	841107	708	.06	*(.06)*	.98	*(.23)*	.11	*(.08)*	1.15	*(.29)*	.515	.387
France	770907	1082	-.02	*(.22)*	1.00	*(.17)*	.08	*(.26)*	.71	*(.16)*	.995	.166
Italy	770907	1082	.18	*(.23)*	1.01	*(.09)*	.52	*(.29)*	.90	*(.21)*	.734	.108
Netherlands	770105	1117	.00	*(.09)*	.95	*(.15)*	-.06	*(.10)*	1.09	*(.25)*	.911	.734
Sweden	770907	1082	.27	*(.14)*	.82	*(.13)*	.19	*(.20)*	.64	*(.16)*	.044	.060
U.K.	770105	1117	.16	*(.16)*	.72	*(.17)*	.43	*(.23)*	1.07	*(.26)*	.048	.157
U.S.	770105	1117	-.04	*(.14)*	1.41	*(.33)*	.03	*(.21)*	.83	*(.20)*	.459	.691

Standard errors have been corrected for heteroskedasticity, and for serial correlation from overlapping observations.

Table II
Estimates and stability tests of the time series process followed by Eurocurrency spreads

$$s_t^6 = a + b s_{t-1}^6 + c[s_{t-1}^{12} - s_{t-1}^6] + \varepsilon_t \qquad \textbf{(26)},$$

and stability tests of the AR(1) subcase.

Country	Start	#obs. (3-mth)	Full sample (1977-98)						1990-98		Subsample stability tests (*P*-values)	
			a	*(s.e.)*	b	*(s.e.)*	c	*(s.e.)*	c	*(s.e.)*	Eq. (26)[a]	AR(1)[b]
Belgium	770914	1081	-.012	*(.016)*	.982	*(.010)*	.037	*(.022)*	.155	*(.082)*	.081	.071
Denmark	770914	1081	-.004	*(.024)*	.963	*(.014)*	**.065**	*(.029)*	.153	*(.100)*	.294	.031
ECU	841114	707	-.019	*(.018)*	.974	*(.015)*	.052	*(.030)*	.061	*(.048)*	.611	.291
France	770914	1081	.001	*(.028)*	.956	*(.042)*	.073	*(.063)*	.042	*(.036)*	.578	.176
Italy	770914	1081	.006	*(.016)*	.989	*(.009)*	.012	*(.020)*	.044	*(.068)*	.061	.063
Netherlands	770112	1116	-.016	*(.038)*	.959	*(.023)*	**.081**	*(.041)*	.042	*(.039)*	.532	.091
Sweden	770914	1081	-.010	*(.010)*	.964	*(.021)*	**.067**	*(.035)*	**.121**	*(.046)*	.365	.148
U.K.	770112	1116	.106	*(.043)*	.955	*(.024)*	.039	*(.046)*	.101	*(.134)*	.085	.061

[a] $\chi^2(3)$ test of stability of $(\hat{a}, \hat{b}, \hat{c})$ estimates across 1977-89 and 1990-98 subsamples,
[b] $\chi^2(2)$ test of stability of $(\hat{a}, \hat{b})$ estimates across 1977-89 and 1990-98 subsamples, with *c* set to zero.
All standard errors and stability tests were adjusted for heteroskedasticity.