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Abstract

Using a simple method, based on forward interest spreads, we analyse the recent movements in the 10-year yield differentials between three currencies (Italian lira; Spanish peseta; Swedish krona) and the DM in order to gauge the extent to which the reduction in these differentials was due to market arbitrage triggered by the expectation of EMU or to more "fundamental" factors (lower inflationary expectations; improved fiscal outlook). We find that most of that reduction cannot be *directly* explained by EMU expectations, though EMU is likely to have had an *indirect* influence by providing an incentive for faster convergence on inflation and fiscal performance. As a by-product of our analysis, we compute estimates of the market probabilities of EMU taking place and of each country joining at different dates (1999 and 2002).

Contents

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1. Introduction: accounting for the recent long-term yield convergence in Europe[1]

Since mid-1995, long-term interest rates in Europe have sharply declined, converging to the lowest levels (Figure 1). Long rate movements during this period followed a pattern similar to that observed in the period following the 1992-93 ERM crisis, a time of record performance of European bond markets.

Two sets of factors are usually invoked to explain this phenomenon. First, domestic economic "fundamentals", in particular inflation, inflationary expectations and fiscal performance, improved significantly in most European countries. Average inflation fell steadily in the EU after 1992 (Table 1). The slowdown of consumer prices was particularly strong for the high-yielders, which broadly coincide with the group countries initially having higher inflation rates. Significant improvements also took place in budgetary performance, partly as a result of fiscal policy strategies directed at meeting the Maastricht Treaty requirements; the average fiscal deficit and its dispersion across countries declined significantly in the period. In particular, significant reductions of the fiscal deficits are estimated to have taken place in 1997, the relevant year for the calculation of the convergence criteria. Better inflation and fiscal performance should have reduced long-term nominal

[1] We thank an anonymous referee for useful comments and suggestions. The usual disclaimer applies.

interest rates in two ways, i.e. by lowering inflationary expectations and narrowing the premia related to market volatility and credit risk.

A second line of explanation centres on the role of financial market trades based on the expectation of EMU. Even if all fundamental economic factors had remained unchanged, the expectation that a single currency would have been introduced at a near future date could have given rise to arbitrage transactions, having the effect of narrowing long-term differentials by reducing (eliminating, in the case of certainty) the forward interest differentials with the DM after that date (a point stressed by De Grauwe, 1996). Several events in 1996 indeed suggested that the introduction of the single currency was within reach. A scenario for the changeover to the single currency had been agreed upon at the December 1995 EU Council in Madrid. In the subsequent months, important steps were made on these issues, namely, the legal status of the euro, the exchange rate system between the euro and other EU currencies and the fiscal rules to be applied in Stage Three. On all three, an agreement was reached during the autumn of 1996 and signed in Dublin in December. A detailed scheme for the operational setting of the European Central Bank was worked out by the EMI in 1996 and published in January 1997.

In principle, the two sets of factors could be identified, i.e. one could try to estimate the impact on long-term rates originating from the improvement in domestic fundamental conditions, for given probability of EMU, and *viceversa*. Such a counterfactual exercise would bring several interesting insights. First, it could shed light on the relevance of the Maastricht convergence criterion based on the long-term interest rate differentials. To the extent that long-term rate convergence reflects "political" developments, unrelated to domestic economic fundamentals, its significance for discriminating among individual countries' performance could be questioned, as was hinted by the Bundesbank (1996, p. 26). Second, evidence that domestic economic fundamentals were, instead, the main driving force behind long-term rate developments would be reassuring for both policy makers and financial markets, because it would mean that market conditions are not excessively sensitive to political uncertainties on EMU and that the convergence of yields will not be easily reversed. In practice, however, the two components are difficult to distinguish, for two main reasons; first, since the efforts aimed at fulfilling the Maastricht criteria are a major factor behind the adjustment policies in many countries, there is a risk of underestimating EMU as the real motive underlying fundamental convergence; second, inflation and fiscal convergence tend to increase a country's likelihood of meeting the criteria, thus increasing any possible "political" effect on long-term rates.

Bearing all these caveats in mind, in this paper we use a simple method, based on forward swap interest spreads, to calculate the relative weight of the two factors and apply it to three countries: Italy, Spain and Sweden. Following what has been recently proposed by other authors,[2] we infer the

[2] De Grauwe (1996), J. P. Morgan (1996), Weidmann (1996), Favero et al. (1997).

market subjective probability of EMU from the post-1999 forward interest spreads with the DM, assuming risk neutrality. However, we depart from these analyses in that we estimate separate EMU probabilities at different horizons after 1999 from the steepness of the forward curve after that date. We thus use these probabilities to decompose the changes in the 10 year yields in the two components, as explained. We focus our analysis on the 1996-97 period, in which the prospects of EMU shifted significantly and, therefore, the related effects on the yield curve may have been important. Our main conclusion is that the decline in long-term yield differentials with the DM on the three currencies considered cannot be attributed mainly to the market's expectation of EMU. This implies that long-term rates should not prove to be excessively sensitive to political uncertainties concerning the EMU process, provided that the improvement in fundamentals, which crucially relies on sound domestic monetary and fiscal policies, is not reversed. Clearly, this conclusion depends on how sensitive domestic monetary and fiscal policies will be to the prospects concerning the realisation of EMU.

The rest of the paper is organised as follows. In Section 2, our estimation method is explained and compared with that used by other authors. In Section 3 we present and discuss our results. In Section 4, we check the robustness of our results by applying a slightly more complex model in which we explicitly distinguish the implications, for each country, of the scenario in which EMU is delayed from that in which it starts on schedule but the country concerned does not immediately join. Finally, Section 5 contains our concluding remarks.

2. A simple model for estimating the impact of EMU on yield spreads

Let Δs^{10} be the change in the 10-year swap interest differential (zero coupon) between a given currency and the DM over a given period of time (the time subscript t can be omitted for brevity). Δs^{10} can be expressed as an average of changes in 1-year forward rate differentials f^j:

$$(1) \qquad \Delta s^{10} = \frac{1}{10} \sum_{j=0}^{9} \Delta f^{j} \; ,$$

where the superscript j denotes the settlement date of the forward contract (Δf^0 corresponds to the change in the spot rate differential). Under the assumption of risk neutrality (and ignoring convexity bias effects[3]), the observed 1-year forward spread coincides with the expected future short rate, which can be written, after January 1st, 1999, as a weighted average of the differentials prevailing in two possible future states:

[3] Forward rates are biased predictors of expected future short rates not only as a result of risk premia but also as a consequence of the non-linearity of the relationship between bond prices and expected future short rates. In particular, due to a Jensen-inequality-type of bias, forward rates tend to overstate future rates to an extent dependent on the magnitude of the volatility of future short rates (see Shiller 1990, pp. 645-47).

(2) $$f^j = p^j r_U^j + (1-p^j) r_N^j \text{ ,}$$

where r_U^j, r_N^j are the realisation of the short rate differential contingent upon, respectively, the participation or non-participation of the given currency in EMU and p^j is the probability of joining EMU at date j. Since under EMU all currency risks should vanish, $(r_U^j = 0)$, the forward spread can be expressed by the last term of equation (2), i.e. the product of the no-EMU interest differential and the corresponding likelihood (we will call no-EMU the case in which the given currency does not participate in EMU in a specified time period). We do not distinguish, at this stage, the case in which EMU does not take place from that in which EMU starts, but the currency does not join; we introduce this distinction in Section 4.

By differentiating equation (2) with respect to t, leaving j constant, we obtain the time changes of the observed forward spread:

(3) $$\Delta f^j = (1-p^j) \Delta r_N^j - \Delta p^j r_{N,-1}^j \text{ ,}$$

where the second term of the right-hand side summarises the impact of the shift in the probability on the observed forward spread changes, while the first term consists in the change in the expected future short spread for the no-EMU case, weighted by the corresponding probability.

Estimating the probability term in equation (2) requires specifying a term structure of forward rates in the no-EMU state. Two approaches have recently been proposed:

1) take a forward spread observed in past periods, when EMU was only a remote possibility (De Grauwe, 1996);

2) estimate a relationship between the spread and a set of explanatory variables expressing world economic and financial conditions, presumably unaffected by EMU; the estimate is performed over a time period when EMU was only a remote possibility and the fitted values of that regression are used to estimate the future no-EMU spread (J. P. Morgan, 1996).[4]

Both approaches share the basic idea that interest rates should return, in the no-EMU scenario, to "normal" conditions observed in the past. The main objection to this procedure is that it implicitly assumes that any change in economic fundamentals occurring recently in the countries concerned, whether induced by EMU-related macroeconomic policies or by other factors, should not be taken into account in estimating the no-EMU spreads after 1999.

In our calculation we use a different assumption, namely, that the probability of a given currency joining EMU can be measured by the *decline* in the forward spread observed

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[4] Weidmann (1996) estimates the no-EMU forward rates of the ecu from the corresponding rates on the component currencies. This estimate, however, is likely to be biased since the forward spreads on individual currencies are themselves affected by the probability of EMU.

for dates from 1999 onwards. In other words, a flat no-EMU term-structure is assumed for the horizons after January 1st, 1999, namely

(4) $r_N^j = r_N^{j-1} \quad \forall j \geq \textit{January } 1° \ 1999$.

In this way, the observed term structure of the forward differential after January 1st, 1999, is assumed to be the weighted average of two values: zero, in the EMU case, and the forward differential prevailing immediately before 1999, in the no-EMU case.

Two basic ideas underlie our very simple identifying assumption (4). The first is that the current term structures of interest rates on European currencies for dates preceding 1999 contain important information on economic policies and prospects in the respective countries, and that this information cannot be disregarded in making projections for more protracted periods. The second is that no compelling reason why, in the no-EMU case, the curve of forward rate spreads vis-à-vis the DM for post-1999 dates should be either upward or downward sloping. Opposite, equally convincing arguments could be made in this respect. On one side, the relaxation of the constraint provided by the goal of meeting the Maastricht criteria could induce, in non-participating countries, a reversal of the orientation of domestic monetary and fiscal policies, leading to a broadening of interest differentials after 1999. On the other, if tight domestic policies have produced long-lasting, favourable effects on market expectations, one could presume that the structure of forward spreads should continue to decline after 1999, independently of participation in EMU. Our assumption of a flat term structure of post-1999 no-EMU forward spreads is neutral with respect to these opposite views.

Based on equation (4), a recursive sequence of implied EMU probability can be recovered by iterating equation (2) at different horizons beyond January 1999:

(5) $$p^j = 1 - \frac{f^j (1 - p^{j-1})}{f^{j-1}}$$

with the understanding that the EMU probability is equal to zero before 1999. Notice that (4) and (5) together imply that $p^j = 1 - \left(f^j / f^{1998}\right)$, where f^{1998} is the forward spread with settlement date 1998. Using (5), equation (3) can be computed for each forward spread entering the 10-year spot yield differential, and the contribution of the probability element in the change of such yield over time can be obtained from equation (1).[5]

The calculations have been performed for three high-yielding currencies: lira, peseta and Swedish krona. Moreover, we repeated the same procedure on the ecu-DM differentials, since, given the automatic conversion clause of the ecu into the euro, this can provide estimates of the probability of

[5] The forward rates have been retrieved from the swap rates using the technique described in Ilmanen (1996).

EMU, q^j.[6] In view of the comparatively high volatility of the ecu forward rates, due to the market imperfections and to the lack of continuously active arbitrage, EMU probabilities were calculated assuming that the ecu forward spreads are affected by a random error, as in the following model:

(6) $$f_t^j = \left(1-q_t^j\right)f_t^{1998} + \zeta_t$$

(7) $$\theta_t^j = \theta_{t-1}^j + n_t$$

(8) $$\theta^j = \log\left(\frac{q^j}{1-q^j}\right)$$

where equation (6) is a stochastic version of (5) and equations (7) and (8) specify a stochastic process for q^j, namely a random walk with a logistic transformation to maintain the probability within the [0-1] interval. The two errors are assumed to have classical properties[7] and to be non correlated. The model was estimated using a Kalman filtering iterative procedure, described in Appendix 1.

In all calculations we assumed that EMU will include the DM. Hence, p^j (probability of a currency joining EMU at time *j*) can be interpreted as a *joint* probability (of EMU starting and of that currency joining it); the corresponding

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[6] This point is noted by De Grauwe (1996).

[7] If f^{1998} is also affected by a random error, equation (6) would be affected by simultaneity bias. For simplicity, we have ignored this complication.

conditional probability can thus be retrieved using the following expression:

$$(9) \qquad p^j = \rho^j q^j$$

where ρ^j is the conditional probability of the country joining EMU at date *j*, given that EMU has started.[8] The estimation was performed using daily time series of forward rates from zero to nine years. In order to match the horizons defined by a set of fixed dates starting on January 1st, 1999, and moving forward at 1-year intervals to 9 years out, a linear interpolation of adjacent 1-year forward rates was used:

$$(10) \qquad f^s = \lambda_{j-s} f^{j-1} + (1-\lambda_{j-s}) f^j \text{ , } s \in [j-1, j], \; s = 1998, 1999, 2000, \ldots$$

where $\lambda_{s-j} \in [0,1]$ is a linear function proportional to the time distance between the chosen set of fixed dates and the set of dates matched by the current 1-year forward term structure.

3. Results and discussion

The 10-year swap rate differential of the three currencies with the DM are shown in Figure 2; the estimated probabilities are shown in Figure 3 (probability of EMU) and 4 to 6 (unconditional probability of each country joining EMU).

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[8] The distinction between joint and conditional probability with regard to EMU participation is stressed by Weidmann (1996).

Two time horizons for all probabilities are considered: 1999 and 2002.

The probability measures can be interpreted in the light of the main institutional and political events during the period. Figure 3 - in which we report both the estimates performed using the deterministic version of (6), and the "smoothed" versions resulting from the full model (6) to (8) - shows that during 1996 the probability of EMU increased steadily, with reference to both 1999 and 2002; the more pronounced rise in the September-December period can be related to the progress made in the institutional preparation for EMU, to which we have referred in a previous section. By early 1997, the EMU seems virtually certain in 1999, and all the more so by 2002. In March, the 1999 probability suddenly declines from over 90 per cent to around 50; this can be related to the increased uncertainty induced by the emerging budget shortfall in Germany and the risk of spillover into Europe of the rise in US official rates. These elements of uncertainty seem to have gradually receded after April, in the wake of the calling and subsequently of the outcome of the French elections. By end-June, the 1999 probability was back to 60 percent; in the subsequent months, it continued to rise up to nearly 90 percent in November, benefitting from the sharp reduction in inflation and the fiscal progress attained in some European countries (e.g. Italy and Spain; see below). It is interesting to note that, during the whole first half of 1997, the 2002 probability level never fell below 95 percent, a sign that the temporary increase in uncertainty during this period did not regard the launch of EMU per se, but only its timing.

In *Italy* (Figure 4), a rising trend in both probability measures is visible after the first quarter of 1996, particularly following the political elections of April 21. A temporary reversal takes place in June, conceivably explained by the announcement by the Government that no attempt would be made to meet the Maastricht deficit criterion in 1997. A sharp rise is visible in September, when the Government presented a 1997 budget law targeting a deficit of 3 per cent of GDP. The probability related to 2002 also increases sharply in the second half of 1996, though following a more volatile pattern. The decline in inflation and survey-based inflationary expectations, particularly strong after mid-1996, is an additional factor explaining the strengthening of market confidence in the second half of the year. By December, in the aftermath of the Dublin agreement and following the approval of the budget law, both probability measures reached a peak. The decline starting in February can be attributed both to the growing market uncertainty on the political feasibility of a broad-based monetary union and to the lower probability of EMU per se, documented in Figure 3. The sharp rise in both probability measures after April 1997 is consistent with the perception that the position taken by the new Socialist Government in France increased the likelihood of a monetary union including many participants from the start. After mid-year, the probability measures of Italy appear to have risen sharply, to reach over 80 percent in November, alongside with a strong reduction of the 12-month inflation rate and a growing confidence that the Maastricht deficit criterion could be met.

It is useful to compare the probability measures computed with our methodology with those of J. P. Morgan's (1996) "EMU-calculator" and with the results of the "EMUPOLL" survey conducted by Reuters; see, in Figure 4a, this comparison for Italy. This comparison suggests a much closer resemblance of our 1999 measure to the Reuters opinion polls, while the "EMU-calculator" tends to mirror our 2002 estimate. This confirms that the J. P. Morgan measure should be better understood as a probability of joining EMU in 1999 or later, rather than in a single year; a point already noted by Favero et al. (1996).

In *Spain*, a sharp, steady upward movement started after mid-1996, in coincidence with the official announcement that it would aim at joining the first wave of EMU founders. In addition, the improved inflation outlook, the stability of the peseta and the consolidation of the recovery during the summer months strengthened market confidence. The unveiling in late September of a restrictive budget (in line with requirements of the convergence criteria) gave a further boost to the chances of Spain. The peak of probabilities is in January, at near 100 per cent for 2002 and 60 per cent for 1999, following the approval of the 1997 budget and the news of decelerating inflation. The sharp downturn in both measures taking place in February, subsequently reversed in April, and the strong recovery in the 1999 probability during the second semester are presumably related to factors similar to the ones already noted for Italy.

Sweden shows a somewhat different and more volatile pattern, largely driven by the official stance towards EMU.

The decline in the first two months of 1996 and the higher level in the second quarter could be explained by domestic political factors, such as the prospect of a referendum on EMU and the fact that a sceptical public opinion appeared to jeopardise the stability of the ruling coalition. A sharp upward trend starts after mid-1996, but only in the probability measure referring to 2002; this can be the result of the wave of optimism on a broad-based EMU prevailing in the second half of the year, coupled with very positive developments on inflation and better prospects for economic recovery. Renewed domestic political uncertainty sharply lowered the probability at the beginning of 1997; its value, in the reference to 1999, was already zero when the Government officially announced the opt-out, in March. On the whole, the probability of Sweden joining EMU in 1999 appears to have been very small in the whole period under consideration. Interestingly, however, the 2002 probability rose in the second half of 1997, which would seem to suggest that the wave of confidence in EMU observed in other European countries may have had a limited effect on Sweden too.

Figures 7 to 9 show the effect of the changes in EMU probabilities on long-term yield differentials with the DM. The Figures show the cumulative change in the differentials relative to the beginning of 1996, and the component attributable to non-EMU factors. The vertical difference between the two lines (shaded area) thus shows the portion that can be attributed to shifts in EMU probabilities at all relevant horizons.

In all cases, the expectation of EMU explains only a small part of the decline in the interest rate differentials. From the beginning of 1996 to November 1997, the portion that can be attributed to it as a percentage of the total decline is 17 per cent for Italy, 22 per cent for Spain; conversely, for Sweden EMU expectation had an opposite effect, reducing the decline in the differential by 13 basis points (the actual decline was 141 basis points). For Spain and Italy most of the effect linked to the increase in EMU probabilities started to cumulate in the fourth quarter of 1996; in Sweden a sharp decline of EMU probability was concentrated in the first quarters of 1996 and 1997. Between end-1996 and November 1997 the component attributed to EMU probability changed only sligthly for the three countries, although with some fluctuation.

One implication of our results is that the temporary increase in the yield spreads occurred in the first quarter of 1997 should be the result of perceived changes in economic fundamentals. Conceivably, one explanatory factor may have been the exchange rate depreciation experienced, in effective terms, in all three countries, due to the strengthening of the US dollar; this could have caused a rebound of inflationary expectations or risk premia. Moreover, the global increase in interest rates triggered by expectations of a monetary policy tightening in the US has probably exerted, as in other instances in the past, an upward pressure on interest rate differentials for high-yielding European countries.

4. Considering the different implications of delaying EMU and of delaying entry into EMU

In the simple model presented in Section 2, we assumed that the no-EMU forward spread would remain constant at its 1998 value, *regardless* of whether the failure to adopt the single currency would result from a delayed entry of the country concerned or to a postponement, or abandonment, of the EMU project as a whole. However, it could be argued that the two cases (which we will call "out-of-EMU" and "no-EMU") may have quite different implications for non-core countries. A country temporarily out of EMU, wanting to fulfil the necessary qualifying criteria for late entry, would probably be subject to strong incentives to maintain strict monetary and fiscal policies, which should help maintain narrow interest margins with core partners that have already adopted the single currency. Conversely, if EMU was delayed, financial markets could perceive an increasing risk of relaxation of domestic macroeconomic policies for the high-yielders. The forward spreads could thus be higher in the no-EMU than in the out-of-EMU scenario.

Retaining the assumption of risk neutrality, we can formalise these ideas by expressing the actual forward spread at horizon j, f^j, in the following way (alternative formulation to equation (2)):[9]

$$(2') \qquad f^j = q^j\left[\rho^j r_U^j + \left(1-\rho^j\right) r_O^j\right] + \left(1-q^j\right) r_N^j$$

[9] This formulation was suggested by a referee.

where ρ^j is the conditional probability defined in equation (9) and r_O^j, r_N^j are the realisations for the future short rate differential in the out-of-EMU and the no-EMU scenarios, respectively. Recalling the definition of conditional probability and that $r_U^j = 0$ after 1999, rearranging terms we can reduce equation (2') to:

$$(11) \qquad f^j = \left(1 - p^j\right) r_O^j + \left(1 - q^j\right)\left(r_N^j - r_O^j\right).$$

Note that if $q^j = 1$ (EMU is certain) or $r_N^j = r_O^j$ (there is no distinction between the no-EMU and the out-of-EMU scenarios) equation (10) reduces to the simpler case considered in Section 2. In all other cases, the estimate of the joint probability will be higher, for given q^j and r_O^j, the higher the no-EMU spread: a higher probability of entry is needed to compensate for a higher spread in the no-EMU case.

Dividing (11) by r_O^j, and rearranging, we obtain:

$$(12) \qquad p^j = \left(1 - \frac{f^j}{r_O^j}\right) + \left(1 - q^j\right) \frac{r_N^j - r_O^j}{r_O^j}.$$

In order to calculate equation (12), we need assumptions for the values of r_N^j and r_O^j, for the relevant time period j, in the distinct no-EMU and out-of EMU scenarios. We have assumed, as in Section 2, that r_O^j is equal to the pre-

1999 forward spread, while we have generated r_N^j by estimating simple reaction functions for the central banks on pre-1996 data,[10] which approximates the monetary policy that would be pursued autonomously by each central bank in the no-EMU case. Monetary policy in the three countries appears to be driven mainly by two variables, inflation and short-term rates in Germany (see Appendix for further details). In simulating the equations, the projected future inflation rates were proxied, in the short-run, by the "Consensus" measures of expected inflation; in the longer run they were assumed to be constant. In a period of rapidly falling inflation like 1996-97, this procedure is likely to yield upwardly biased estimates of expected inflation and, consequently, of future interest differentials in the no-EMU case. In fact, simulated r_N^j values turn out to be significantly above r_O^j in all cases; consequently, the new estimates of the joint probabilities (we will call them "adjusted" probabilities) are systematically larger than the ones computed in Section 2, by an amount given by the second term with the right hand side of (11). In light of the previous considerations, we regard this as an upper bound to the true probability values.

The results are shown in Figures 10-12 (with reference to the 1996 to mid-1997 period), whereas Table 2 illustrates the results of the decomposition of the 10-year yields, obtained with the "adjusted" probabilities. The Figures show that the adjusted estimation produces no significant changes

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[10] A similar approach is followed by Favero et al. (1997).

on the probabilities for 2002, while the results relative to 1999 are more mixed. For Italy, the old and the adjusted measures differ, on average, by 13 percentage points; at end-June 1997, the adjusted probability is around 60 per cent, against 40 per cent for the old measure. For Spain the gap is wider, around 26 percentage points; it approaches 30 per cent at the end of the period, when the adjusted probability nears 75 per cent. For Sweden the adjusted calculation yields implausible results, with the 1999 probability being higher than the 2002 measure.

What is more important for our purpose is that the decomposition of the 10-year yield differentials into EMU-related and other factors is not radically altered by substituting of the adjusted probabilities for the unadjusted ones. As-shown in Table 2, the portion that is attributed to EMU rises from 19 to 36 for Italy and from 30 to 36 for Spain. A more significant change occurs for Sweden, where the EMU-related component increases from a small negative value to almost 50 per cent; however, we are inclined to discount this result, in light of the implausible values assumed, for Sweden, by the adjusted probabilities.

5. Conclusions

Our analysis suggests that the improvement in domestic economic fundamentals is the main factor behind the strong reduction in long-term interest differentials with the DM observed in 1996 and 1997 in Italy, Spain and Sweden. Changes in the market-perceived probability of these countries joining

EMU, which we estimate using a method based on forward interest spreads, explain a much smaller part. However, we cannot exclude that EMU may have had an indirect influence, by fostering inflation and fiscal convergence. The main implication is that the convergence of interest rates should not be overly sensitive to shifts in the political sentiments concerning EMU and its timing, provided that the recent improvement in economic fundamentals, achieved mainly through tight monetary and fiscal policies, is not reversed.

APPENDIX

1. A Kalman filtering procedure for computing smoothed EMU probability

The procedure requires the trasformation of model (6)-(9) into the canonical state-space representation

(1a) $$f_t^j = \left(1/1+\exp(\theta_t^j)\right) f_t^{1998} + \zeta_t \qquad \zeta_t \triangleright N(0,\sigma_\zeta)$$

(2a) $$\theta_t^j = \theta_{t-1}^j + n_t \qquad n_t \triangleright \mathrm{N}(0,\sigma_\eta).$$

Model (1a)-(2a), which is functionally non-linear, only allows an approximately optimal filter by using the extended Kalman procedure, as oulined in Harvey (1989, Section 3.7.2, pp. 160-62). The non-linearity in the measurement equation (1a) can be dealt with by expanding the right-hand-side in Taylor series around the conditional expected value of the state variable, $\hat{\theta}_{t|t-1}^j$:

$$f_t^j = \hat{z}_{t|t-1}\theta_t^j f_t^{1998} + \hat{a}_{t|t-1} + \zeta_t$$

(3a) $$\text{where: } \hat{z}_{t|t-1} = -\frac{exp\left(q_{t|t-1}^j\right)}{1+exp\left(q_{t|t-1}^j\right)}$$

$$\hat{a}_{t|t-1} = \left(\hat{z}_{t|t-1}\theta_t^j + \frac{1}{1+exp\left(\hat{\theta}_{t|t-1}^j\right)}\right) f_t^{1998}.$$

The estimated state variable, θ_t, can be obtained by solving the modified Kalman filter recursions obtained via the linearised state-space model (2a)-(3a):

$$
\begin{aligned}
&\theta_t^j = \hat{\theta}_{t|t-1}^j + k_t(f_t^j - \hat{f}_t^j) \\
&\hat{f}_t^j = \frac{f_t^{1998}}{1+\exp\left(\hat{\theta}_{t|t-1}^j\right)} \\
&k_t = \frac{V_{t|t-1}\hat{Z}_{t|t-1}f_t^{1998}}{1+V_{t|t-1}\left(\hat{Z}_{t|t-1}f_t^{1998}\right)^2} \\
&V_{t+1|t} = V_{t|t-1} - \frac{\left(V_{t|t-1}\hat{Z}_{t|t-1}f_t^{1998}\right)^2}{1+V_{t|t-1}\left(\hat{Z}_{t|t-1}f_t^{1998}\right)^2} + \omega \\
&\hat{\theta}_{t|t-1}^j = \theta_{t-1}^j \\
&\omega = \frac{\sigma_\zeta}{\sigma_\eta}
\end{aligned}
\tag{4a}
$$

where ω represents the signal-to-noise ratio and $V_{t|t-1}$ the conditional variance of the state variable, θ_t. For these latter variables, initial conditions are set to 1, 1 and 0, respectively; the first 50 observations of the recursion are discarded, in order to properly calibrate the initial conditions. The estimated smoothed probability, $q_{t|t-1}$, is thus obtained by inverting equation (8) and computing:

$$
q_{t|t-1}^j = \frac{\exp\left(\hat{\theta}_{t|t-1}^j\right)}{1+\exp\left(\hat{\theta}_{t|t-1}^j\right)}
$$

for the sample period January 1996-June 1997.

2. Reaction functions estimation

We estimate the central bank reaction function for Italy, Spain and Sweden, based on the following very simple specification:

$$\text{(5a)} \quad i_t = \alpha_0 + \alpha_1 i_{t-1} + \alpha_2 \pi_t + \alpha_3 i_t^{ger}$$

where i_t is the end-of period 3-month (LIBOR) interest rate, π_t denotes inflation rate (change in the logarithm of CPI in period t) and i_t^{ger} the end-of-period 3-month Euro-DM rate. Other, more complex specifications including cyclical factors and measures of fiscal deficits did not significantly improve the fit of the equations, so we preferred to work with a more parsimonious, though extremely simple, formulation.

Parameters were estimated on quarterly data covering the sample period 1980-1995 and are reported in Table 3. Statistical tests share a fairly substantial cross-country similarity; in addition, parameter values are similar across countries. For each country estimated future interest rates for the horizon considered, e.g. 1999 and 2002, are projected by replacing the future German rates with the corresponding current forward rates, while expected inflation rates are taken from the "Consensus" forecasts. Since "Consensus" projections only relate to current and next year inflation, periods not covered by the survey were assumed to be constant and equal to the last "Consensus" value. In the simulation, the "Consensus" projections are revised quarterly; each quarter's projection is then linearly interpolated to generate a daily time-series of interest rates.

Table 1

Inflation, Deficit and Long-Term Interest Rates in the EU (1)
(in percent)

Country \ Year	1992	1993	1994	1995	1996	1997(2)	1992	1993	1994	1995	1996	1997(2)
	CPI Inflation % (3)						General Govt. Surplus (+) or Deficit (-)					
Belgium	2.4	2.8	2.4	1.5	1.8	1.7	-7.2	-7.5	-5.1	-4.1	-3.3	-2.6
Denmark	2.1	1.3	2.0	2.1	1.9	2.2	-2.9	-3.9	-3.5	-1.6	-1.4	1.3
Germany	4.0	3.6	2.7	1.8	1.2	1.8	-2.8	-3.5	-2.4	-3.5	-4.0	-3.0
Greece	15.9	14.4	10.9	9.3	7.9	5.7	-12.3	-14.2	-12.1	-9.1	-7.9	-4.2
Spain	5.9	4.6	4.7	4.7	3.6	2.0	-3.6	-6.8	-6.3	-6.6	-4.4	-2.9
France	2.4	2.1	1.7	1.7	2.1	1.3	-3.8	-5.6	-5.6	-4.8	-4.0	-3.1
Ireland	3.1	1.5	2.4	2.5	1.7	1.5	-2.5	-2.4	-1.7	-2.0	-1.6	0.6
Italy	5.4	4.2	3.9	5.4	4.0	1.9	-9.5	-9.6	-9.0	-7.1	-6.6	-3.0
Luxembourg	3.2	3.6	2.2	1.9	1.2	1.6	0.8	1.7	2.6	1.5	0.9	1.6
The Netherlands	3.2	2.6	2.7	2.0	1.5	2.1	-3.9	-3.2	-3.4	-4.0	-2.6	-2.1
Austria	4.1	3.6	3.0	2.2	1.8	1.5	-1.9	-4.2	-4.4	-5.9	-4.3	-2.8
Portugal	8.9	6.5	5.2	4.1	2.9	2.2	-3.6	-6.9	-5.8	-5.1	-4.0	-2.7
Finland	2.9	2.2	1.1	1.0	1.5	1.2	-5.9	-8.0	-6.2	-5.2	-3.3	-1.4
Sweden	2.6	4.7	2.3	2.9	0.8	0.9	-7.8	-12.3	-10.8	-8.1	-3.9	-1.9
United Kingdom	4.7	3.0	2.4	2.8	2.9	2.9	-6.3	-7.8	-6.8	-5.8	-4.6	-2.0
Memorandum item:												
Mean (unweighted)	4.7	4.0	3.3	3.1	2.5	2.0	-4.9	-6.3	-5.4	-4.8	-3.7	-1.9
St. Dev. (unweighted)	3.6	3.2	2.4	2.1	1.8	1.1	3.3	4.0	3.6	2.7	2.1	1.7

(1) Source: EU Commission (April 1997). (2) EU Commission Forecast (Autumn 1997). (3) For 1996-97 EU harmonised CPI.

Table 2

BREAKDOWN OF THE CHANGE IN THE 10 YEAR YIELD DIFFERENTIAL WITH EURO-DM SWAPS BETWEEN DECEMBER 1995-JUNE 1996

(basis points)

Country	Component attributable to:		
Based on: unadjusted probability	EMU	Other Factors	Total
Italy	-56	-246	-302
Spain	-83	-192	-275
Sweden	16	-134	-118
Adjusted probability			
Italy	-108	-194	-302
Spain	-95	-180	-275
Sweden	-56	-62	-118

Table 3

CENTRAL BANK REACTION FUNCTION ESTIMATION

(quarterly data; t-ratios in parentheses)

Dependent variable: 3-month rate	Italy	Spain	Sweden
Coefficient:			
Constant	1.55 (1.23)	0.86 (1.04)	1.81 (2.31)
Inflation (1)	0.19 (2.06)	0.23 (2.26)	0.09 (2.16)
3-month DM rate	0.27 (1.83)	0.10 (1.03)	0.21 (1.93)
Lagged 3-month rate	0.65 (6.83)	0.76 (9.23)	0.66 (7.22)
R^2 corrected	0.78	0.71	0.72
DW (H-statistics)	-1.72	0.13	0.24
Standard error of regression	2.53	1.70	1.69
N° of observations	64	64	64

(1) Quarterly annualized rate for Italy and Sweden; 12-month rate for Spain.

Figure 1

10-YEAR SWAP RATES IN SELECTED EU COUNTRIES

(weekly data; in percent)

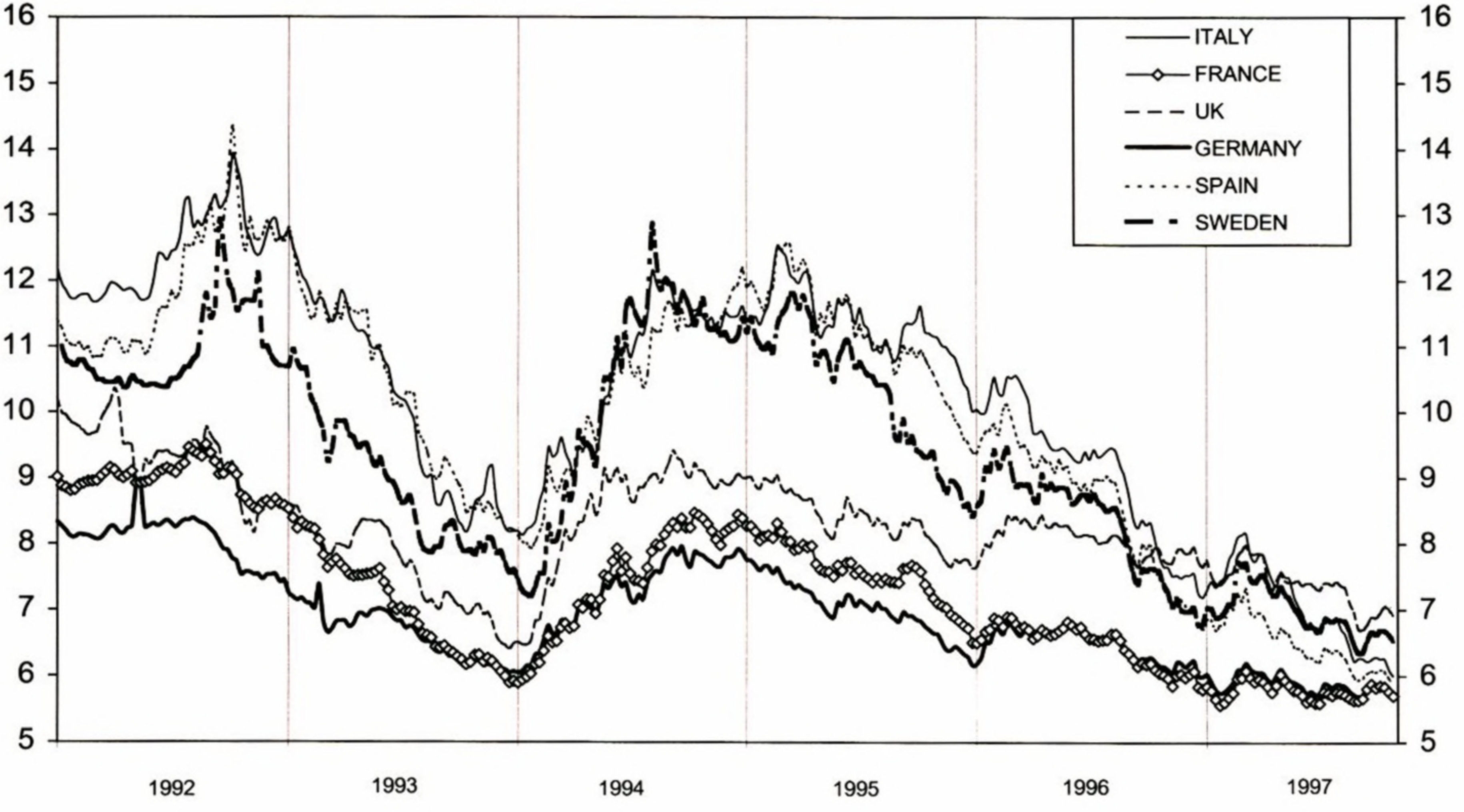


Figure 2

10-YEAR SWAP RATE DIFFERENTIAL WITH GERMANY: ITALY, SPAIN, SWEDEN

(weekly data; in percent)

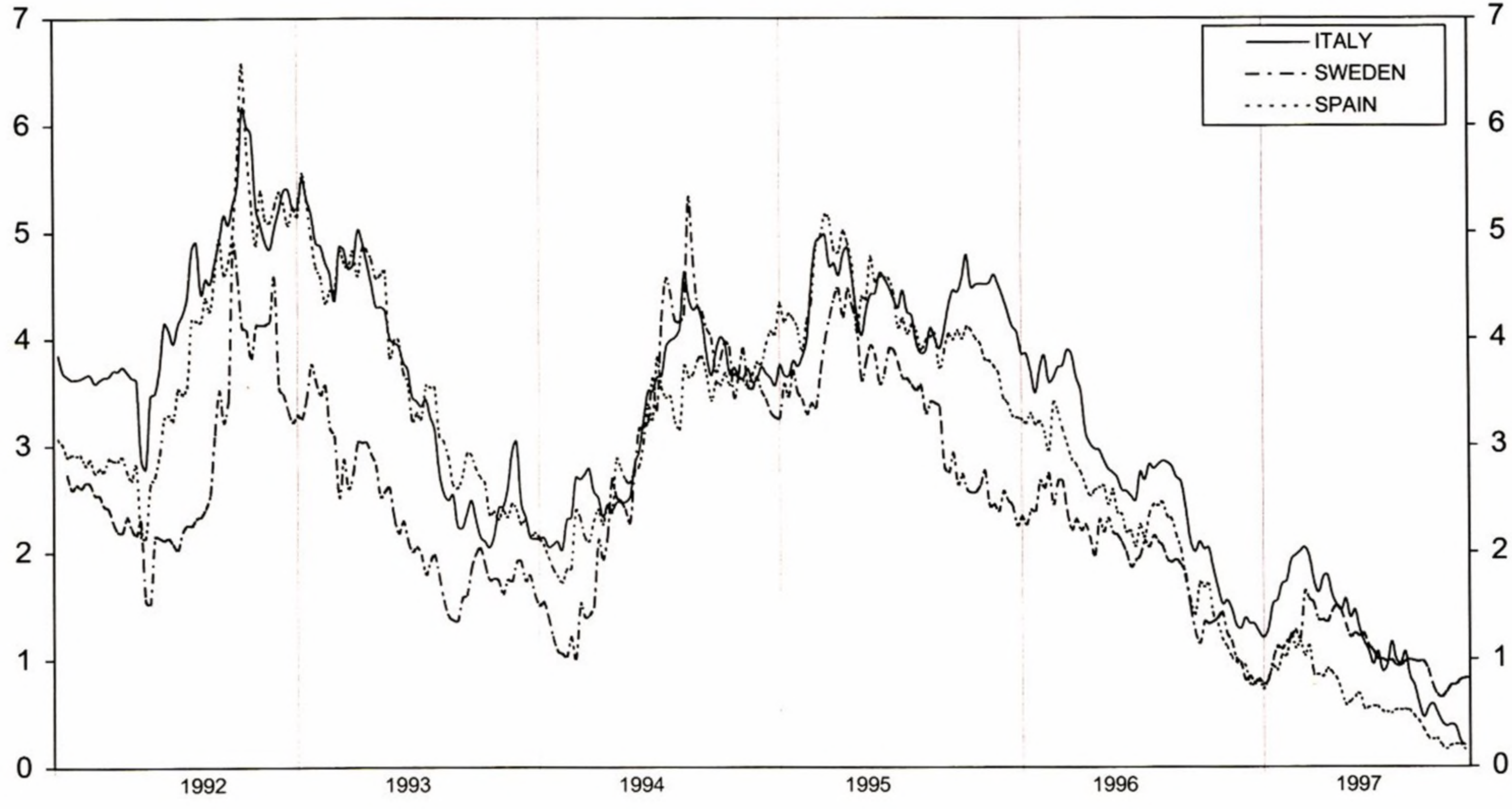


Figure 3

PROBABILITY OF EMU STARTING BY 1999 AND 2002

(daily data; in percent)

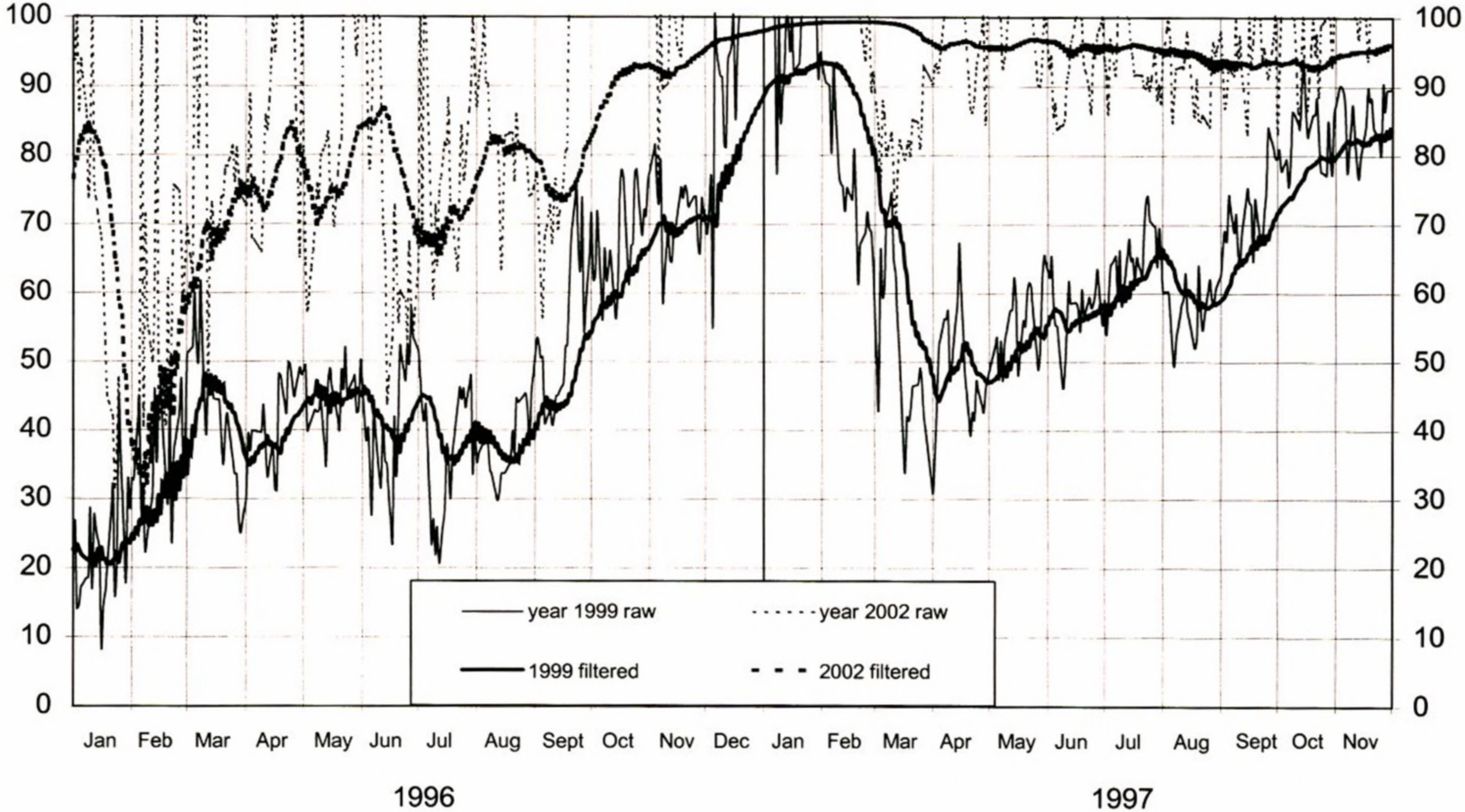


Figure 4

PROBABILITY OF ITALY BELONGING TO EMU BY 1999 AND 2002

(daily data; in percent)

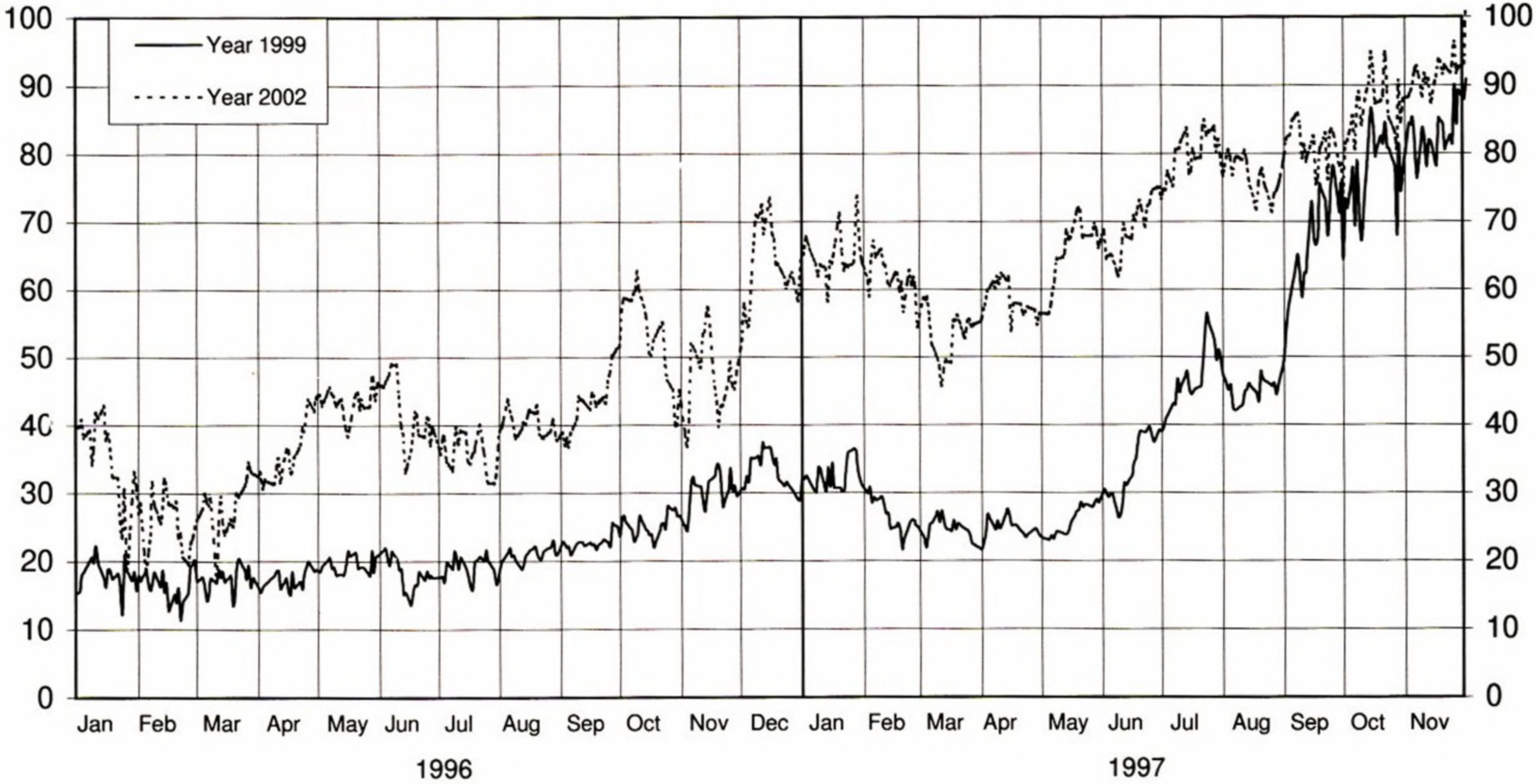


Figure 4a

PROBABILITY OF ITALY BELONGING TO EMU BY 1999 AND 2002

(daily data; in percent)

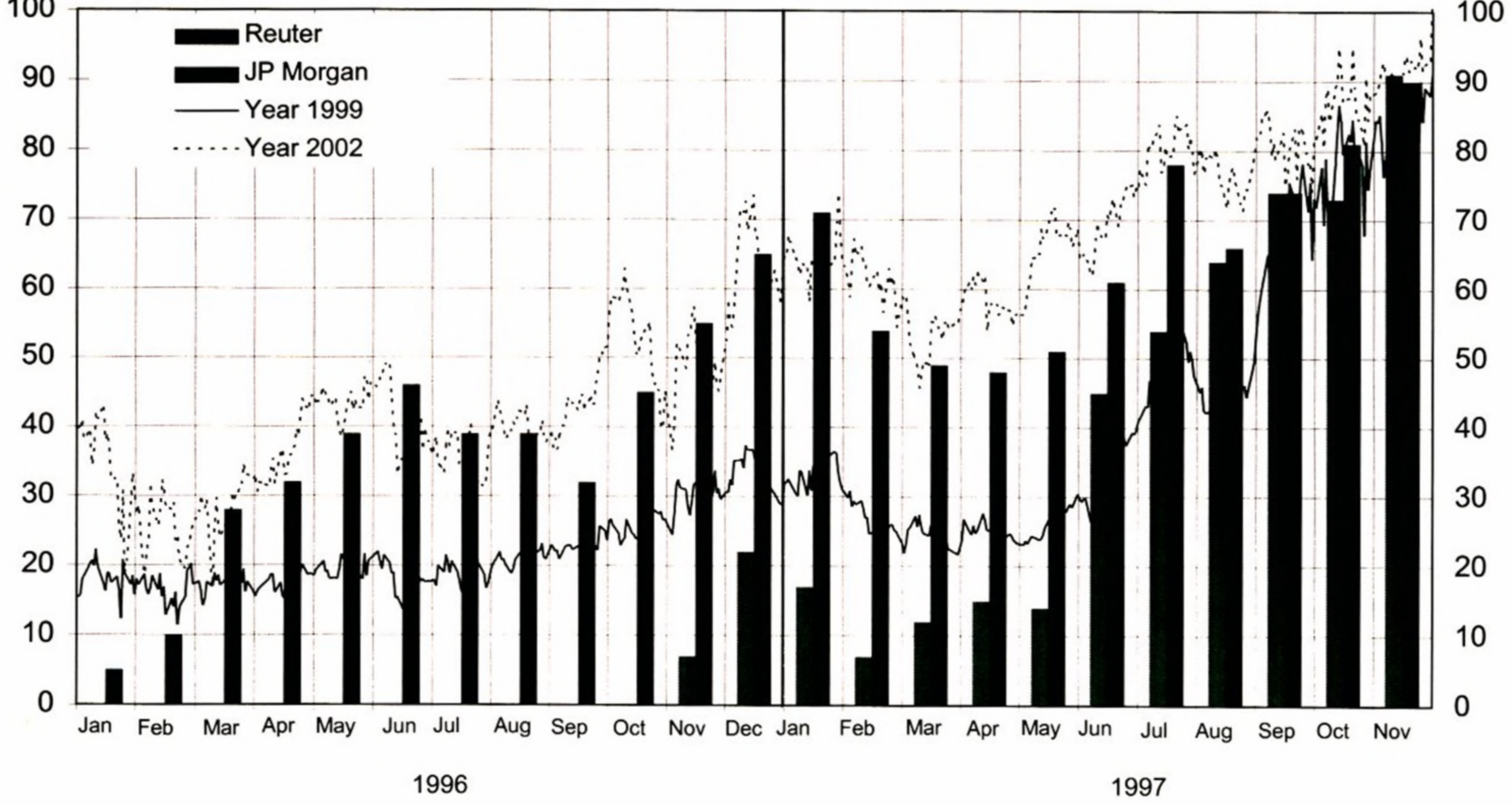


Figure 5

PROBABILITY OF SPAIN BELONGING TO EMU BY 1999 AND 2002

(daily data; in percent)

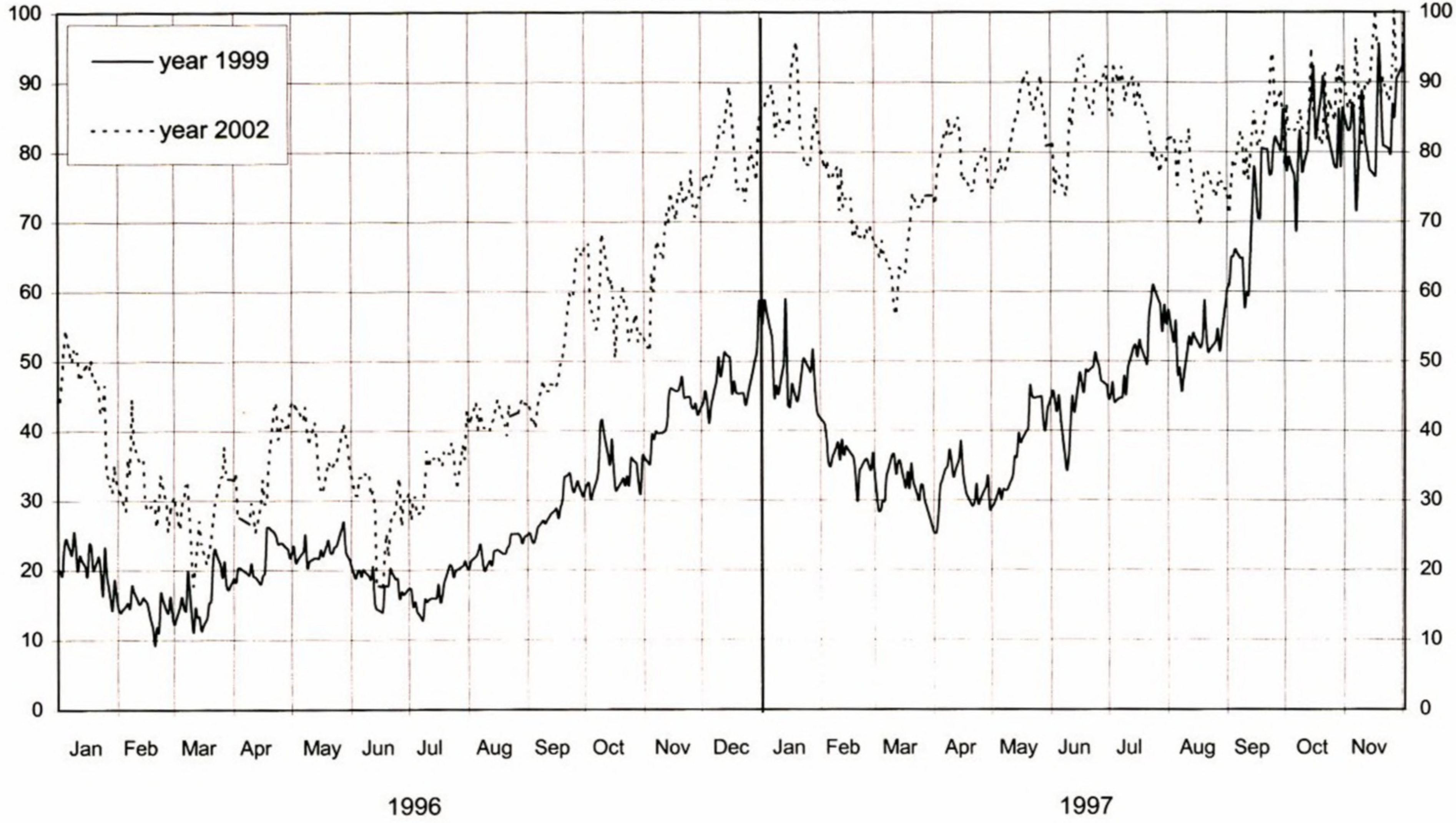


Figure 6

PROBABILITY OF SWEDEN BELONGING TO EMU BY 1999 AND 2002

(daily data; in percent)

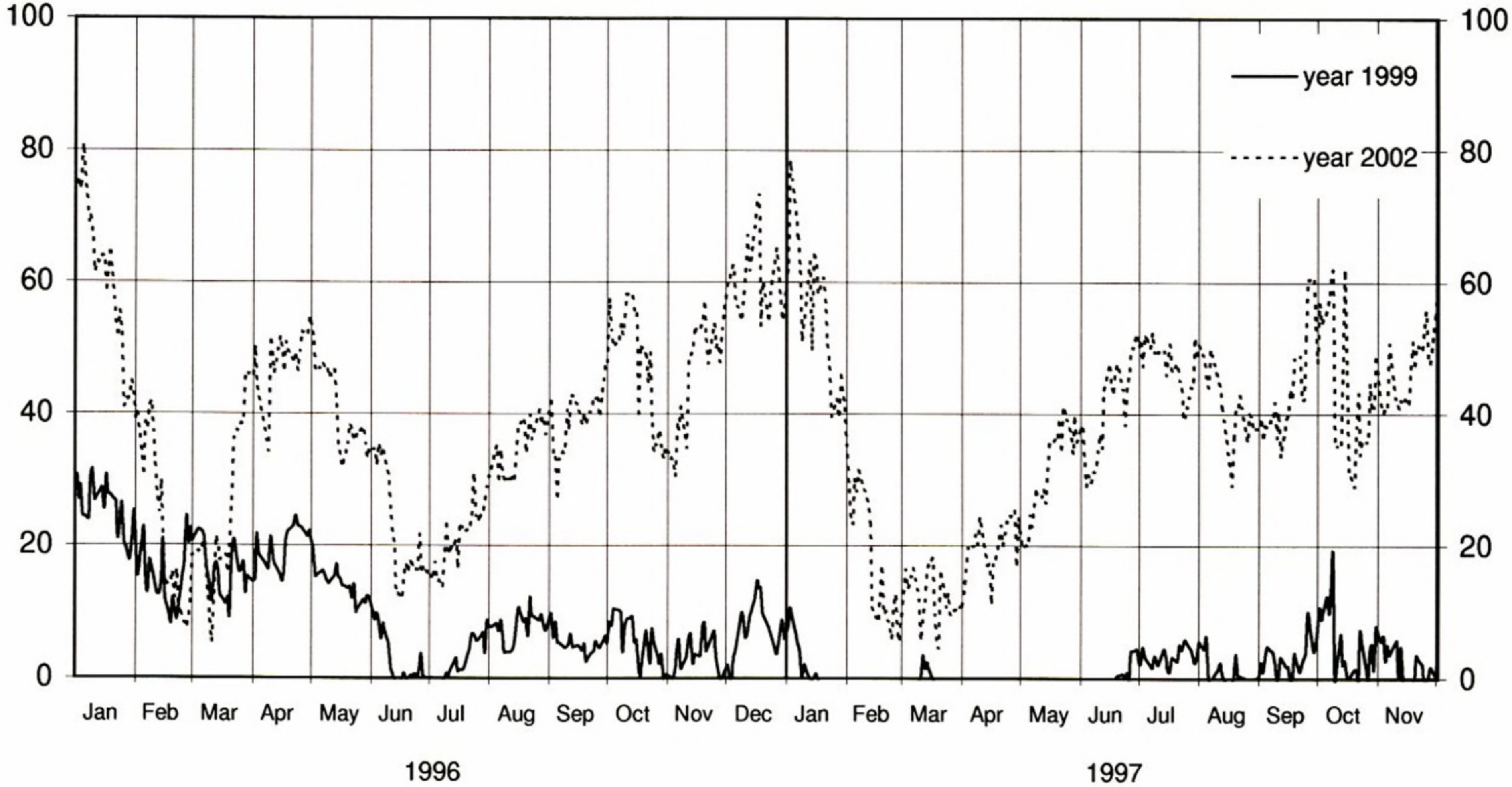


Figure 7

DECOMPOSITION OF THE CHANGE IN 10-YEAR SWAP RATE DIFFERENTIAL WITH GERMANY: ITALY

(daily data; in percent)

Figure 8

DECOMPOSITION OF THE CHANGE IN 10-YEAR SWAP RATE DIFFERENTIAL WITH GERMANY: SPAIN

(daily data; in percent)

Figure 9

DECOMPOSITION OF THE CHANGE IN 10-YEAR SWAP RATE DIFFERENTIAL WITH GERMANY: SWEDEN

(daily data; in percent)

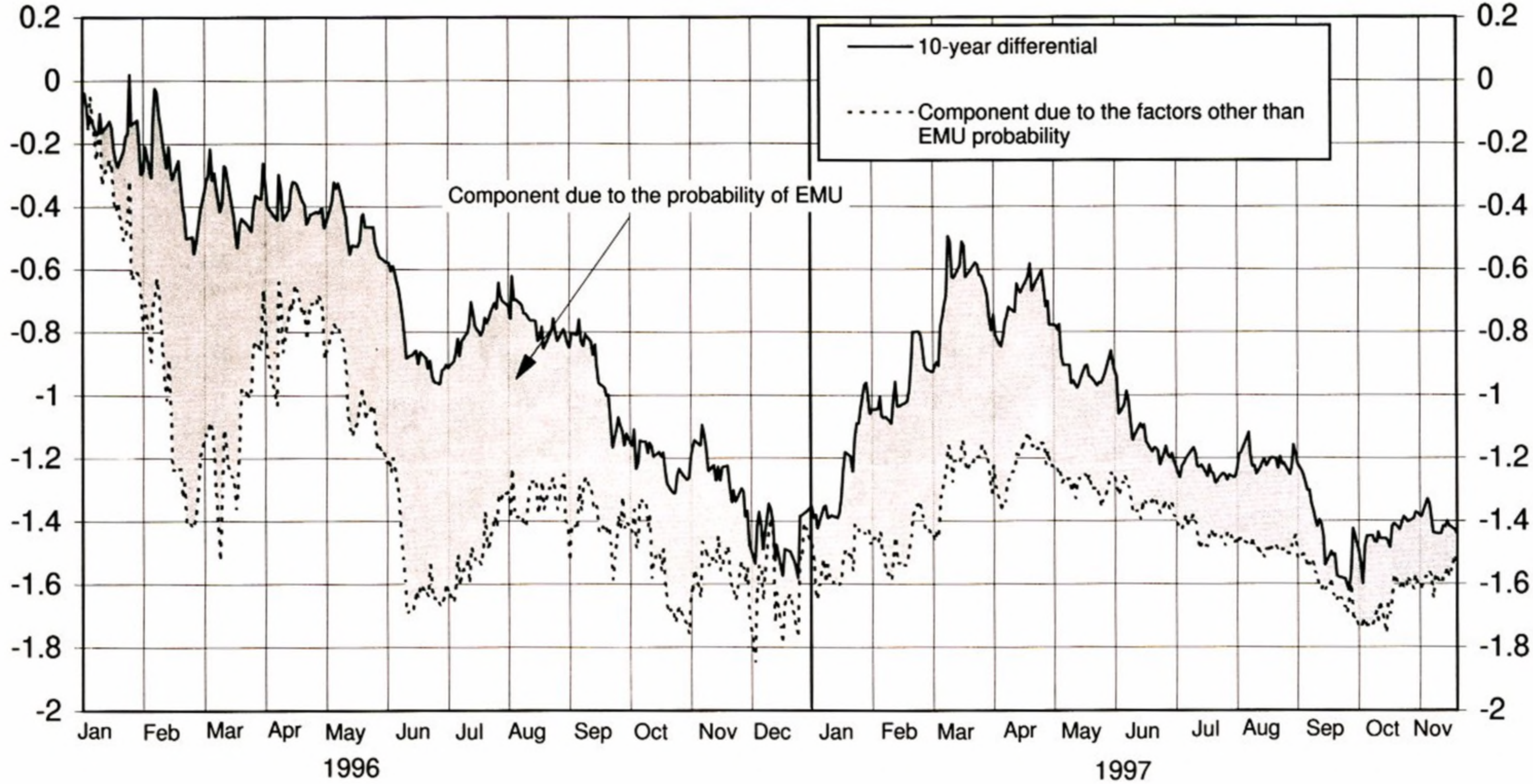


Figure 10a

PROBABILITY OF ITALY BELONGING TO EMU IN 1999

(daily data; in per cent)

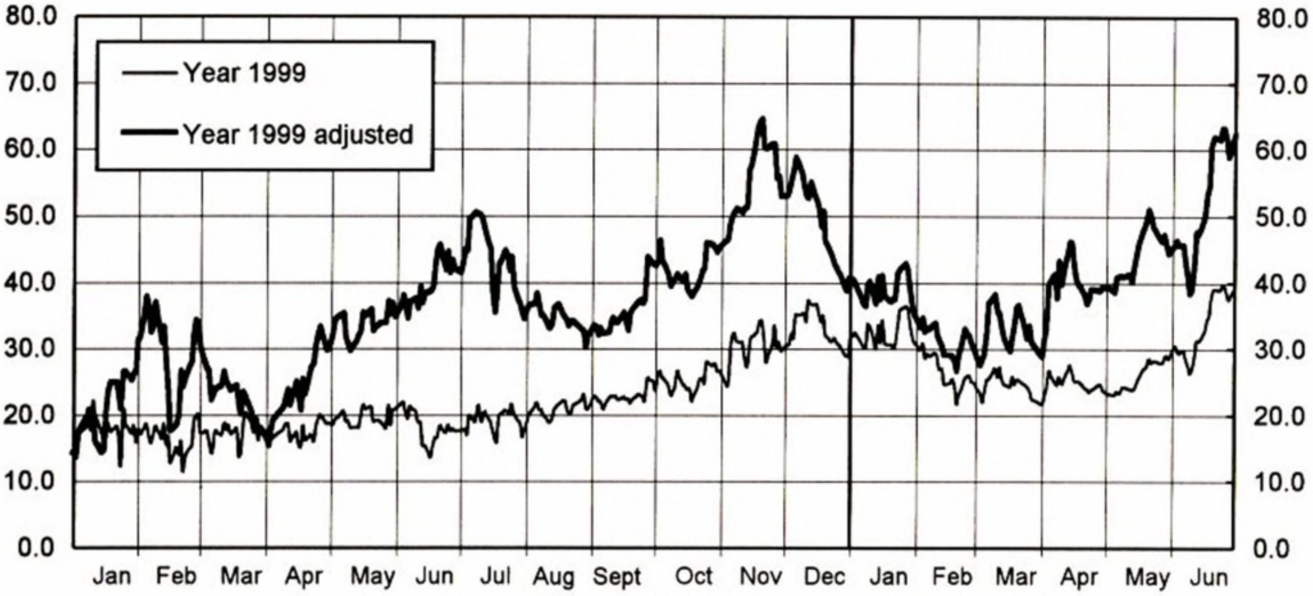


Figure 10b

PROBABILITY OF ITALY BELONGING TO EMU IN 2002

(daily data; in per cent)

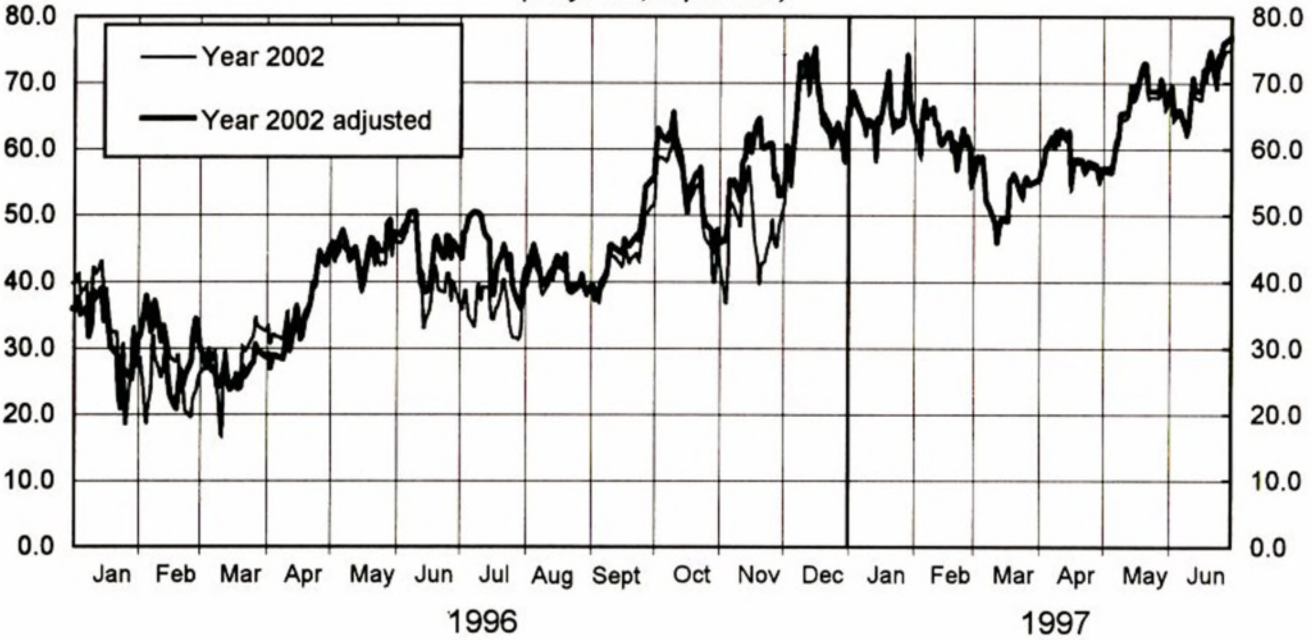


Figure 11a

PROBABILITY OF SPAIN BELONGING TO EMU IN 1999

(daily data; in per cent)

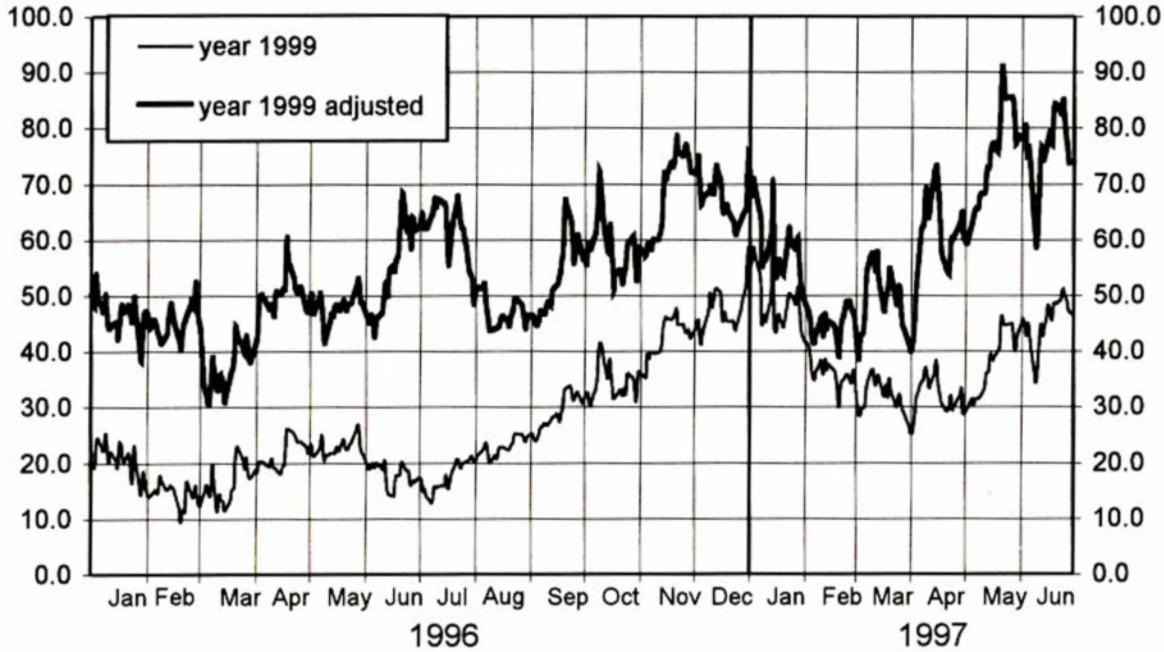


Figure 11b

PROBABILITY OF SPAIN BELONGING TO EMU IN 2002

(daily data; in per cent)

Figure 12a

PROBABILITY OF SWEDEN BELONGING TO EMU IN 1999

(daily data; in per cent)

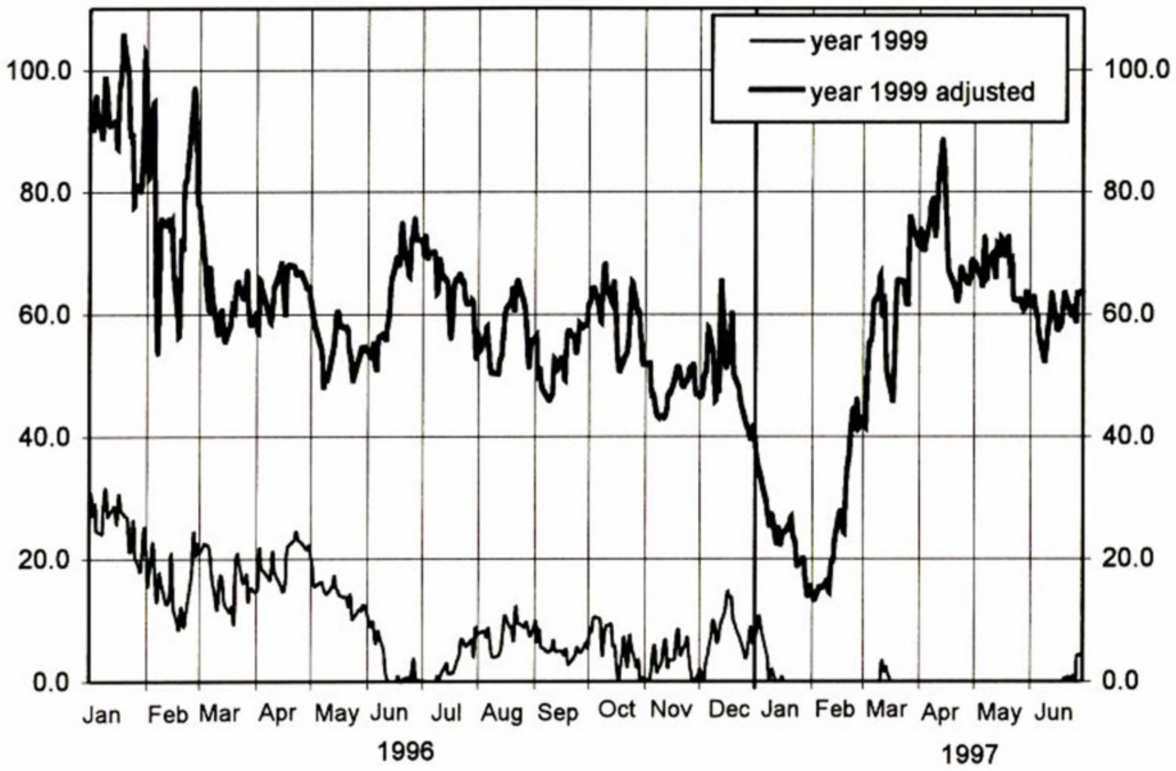
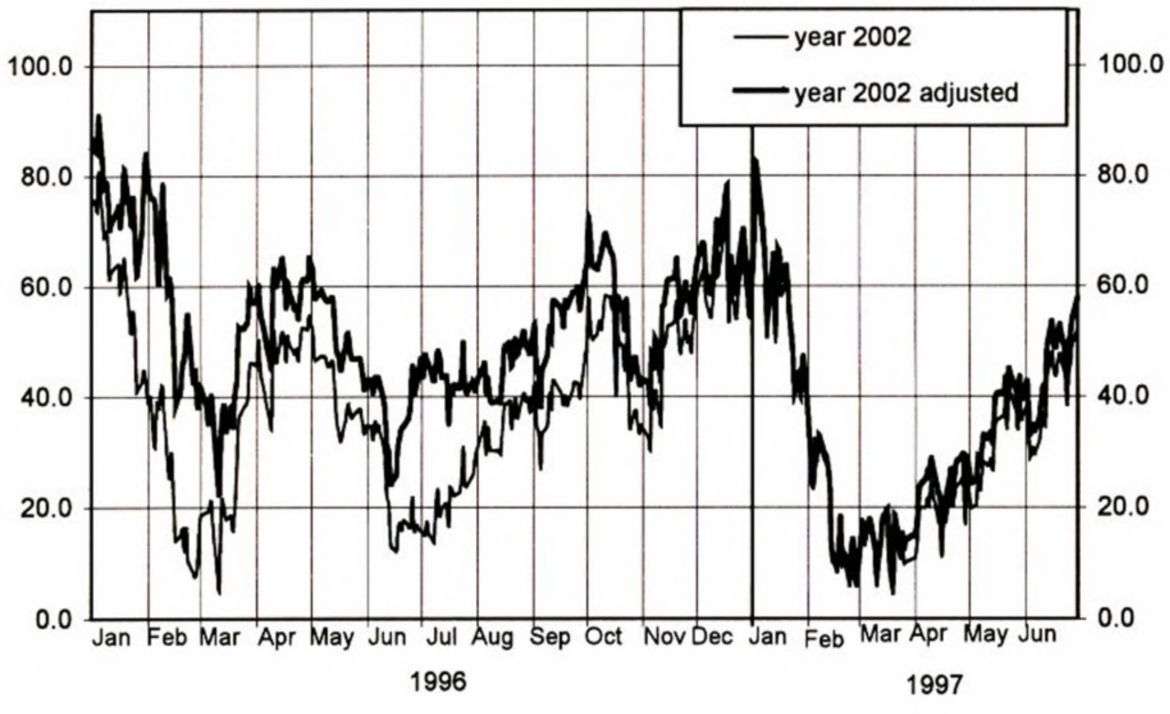


Figure 12b

PROBABILITY OF SWEDEN BELONGING TO EMU IN 2002

(daily data; in per cent)



References

Banco de España (1996), *Economic Bulletin*, October.

Bundesbank (1996), *Monthly Bulletin*, November.

De Grauwe, P. (1996), *Forward Interest Rates as Predictors of EMU*, Centre for Economic Policy Research, Working Paper, No. 1395.

Favero, C. A., F. Giavazzi, F. Iacone and G. Tabellini (1997), *Extracting Information from Asset Prices: The Methodology of EMU Calculators*, mimeo.

Harvey, A. C. (1989), *Forecasting Structural Time Series Models and the Kalman Filter*, Cambridge University Press, Cambridge.

Ilmanen, A. (1996), *Market Rate Expectations and Forward Rates*, *Journal of Fixed Income*, Vol. 6, No. 2.

J. P. Morgan (1996), *EMU Calculator Workbook*, New York.

Shiller, R. J. (1990), *The Term Structure of Interest Rates*, in B. M. Friedman and F. H. Hahn (eds.), *Handbook of Monetary Economics*, Elsevier, Amsterdam.

Sveriges Riksbank (1996), *Quarterly Review*, 4th Quarter.

Weidmann, J. (1996), *The Likelihood of European Monetary Union*, *BNL Quarterly Review*, Vol. 49, No. 199.

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