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FEDERAL RESERVE BANK OF CLEVELAND

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Keywords: Exchange Rates, Central Bank Intervention, Risk Premium, FIGARCH.

JEL Classification numbers: C22, E41, E31

CENTRAL BANK INTERVENTION AND
OVERNIGHT UNCOVERED INTEREST RATE PARITY

by

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Abstract

This paper considers the impact of U.S. and German central bank intervention on the risk premium in forward foreign exchange markets. The model estimation is facilitated with the use of daily data on overnight Eurocurrency deposit rates, so that the interest rate maturity time of one day matches the sampling interval of the data. We also use the official net daily purchases and sales of dollars vis-à-vis the German Mark by the Federal Reserve System and the Bundesbank. The model involves FIGARCH innovations to model the degree of long term dependence in the volatility process. Some support is found for the intervention variables affecting the risk premium as predicted by theory. The impact of intervention in the two years immediately following the meltdown of the equity markets in October 1987 is particularly strong.

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1. Introduction

While central bank intervention has at times been quite substantial in the post Bretton Woods era, there continues to be controversy over its effectiveness in achieving the policy goals of either changing the level of nominal exchange rates or of reducing volatility. A large literature has examined the usefulness of intervention; see Edison (1993) and Almekinders (1995). The general conclusion seems to be that the policy either has consequences that vary with the sample period, effects that are inconsistent with the theory, or ultimately has little impact on nominal exchange rates. On the one hand, Dominguez and Frankel (1993) and Ghosh (1992) find support for the effectiveness of intervention. However, Baillie and Osterberg (1997a, 1997b) find evidence that G3 intervention either has no statistical effect, or that it has outcomes which are the opposite to those intended.

Econometric analysis of the impact of central bank intervention is generally constrained by the availability of official data. Clearly the nature of the volatility of asset price markets and the likely short-lived nature of risk premia make it desirable to use daily data or very high frequency data. Ideally, we would utilize intraday data, as in the study by Goodhart and Hesse (1993). In this study we use the officially recorded net intervention by the Federal Reserve Bank over the previous 24 hours, and similar but confidential data kindly supplied by the Bundesbank.

The appropriate sampling frequency of the data is related to the hypothesized transmission mechanism of intervention. Studies such as Ghosh (1992), Obstfeld (1989) and Humpage (1988) which relate intervention to portfolio balances are limited to monthly data and have been generally inconclusive as to the validity of the portfolio balance effect. The possibility that intervention signals a change in monetary policy is examined by Dominguez (1990) who uses weekly data on monetary surprises, exchange rates and intervention and finds its effectiveness varies with the credibility of monetary policy. Conversely, if an 'small' intervention can provide a signal of monetary policy (Klein (1992)), then the impact of the signal might be largely unrelated to fundamentals such as money supply.

Generally, studies that have examined the effects of daily intervention on daily spot exchange rates have either found no effects (e.g. Baillie and Osterberg (1997a)), or effects that are extremely weak, e.g. Goodhart and Hesse (1993). A small number of studies have assessed intervention's impact on the mean and/or conditional variance of deviations from uncovered interest rate parity (UIRP). Loopesko's (1984) approach utilized cumulative intervention flows from the beginning of her sample period while Dominguez (1992) analyzed the impact of daily flows. Humpage and Osterberg (1992) tried both approaches in their analysis of the conditional mean and conditional variance of deviations from UIRP.

One theoretical motivation for an impact of intervention on UIRP is provided by Osterberg (1989) and Baillie and Osterberg (1997b) who formulated a two country inter-temporal asset pricing model which implied that central bank foreign exchange intervention affects the forward exchange risk premium. Baillie and Osterberg (1997b) found empirical support for intervention influencing the risk premium in the forward DM-$ and Yen-$ markets. Purchases of dollars by the Federal Reserve System were found to be associated with excess $ denominated returns, and furthermore, there was evidence that intervention increased rather than reduced exchange rate volatility.

The Role of Intervention

The Federal Reserve System and the Bundesbank appear to routinely sterilize their interventions so that the purchase (sale) of foreign currency is offset by a corresponding sale (purchase) of domestic government debt to eliminate the effects on domestic money supply. Clearly, unsterilized intervention is equivalent to monetary policy and is more likely to directly effect exchange rates. However, even sterilized intervention might be linked to monetary policy.

The literature views intervention as working either by signaling the central bank's future monetary policy or by operating via a portfolio-balance effect. The latter approach is motivated by mean-variance optimization, where agents are concerned with terminal wealth composed of domestic and foreign currencies and bonds. Sterilized intervention will alter the relative supplies of domestic money and bonds. With risk averse investors who view domestic and foreign bonds as imperfect substitutes, the impact of intervention will adjust the relative rate of return by changing the exchange rate. However, the portfolio balance theory implies no impact of intervention on the exchange rate when there is perfect substitutability of bonds and/or Ricardian equivalence, so that consumers exactly anticipate future taxes associated with government debt. Any test of the theory requires information on the relative supplies of the assets.

The alternative view of intervention as a signal of the central bank's future monetary policy implies that a sterilized purchase of foreign currency is expected to lead to a depreciation of the exchange rate if the foreign currency purchase is assumed to signal a more expansionary domestic monetary policy. Klein and Rosengren (1991) find no consistent relationship between intervention and monetary policy and Kaminsky and Lewis (1996) report that the impact of intervention on exchange rates has sometimes been inconsistent with the implied monetary policy. Humpage (1997) concludes that the US authorities in the 1990s had no information superior to the market so that intervention could not be viewed as signaling new information about monetary policy. Dominguez and Frankel (1993) found inconclusive evidence on the signalling and portfolio balance transmission mechanisms, while Ghosh (1992) found that variables associated with the portfolio balance approach appear to have little effect on spot exchange rates.

Definitions

For subsequent analysis, we define S_t as spot exchange rate in terms of DM-$ at time t, F_t is the forward exchange rate at time t, for delivery at time t+k and P_t is the domestic price level. The UIRP condition is,

$$(1) \qquad E_t[(F_t - S_{t+k})/P_{t+k}] = 0,$$

since expected real returns in the forward market are zero. On taking a Taylor series expansion to second order terms,

$$(2) \qquad E_t s_{t+k} - f_t = -(1/2)Var_t(s_{t+1}) + Cov_t(s_{t+1}p_{t+1}),$$

$$(3) \qquad E_t \Delta s_{t+1} - (i_t^* - i_t) = -(1/2)Var_t(s_{t+1}) + Cov_t(s_{t+1}p_{t+1}),$$

where lower case variables denote the logarithms of variables in levels, and i_t is the dollar return on a risk free $ denominated bond, and i_t^* is the foreign currency return on a risk free bond denominated in terms of the foreign currency. Usually, the ex post deviation from UIRP is expressed as, $y_t = \Delta s_{t+k} - (i_t^* - i_t)$, where the two terms on the right hand side of (2) are neglected. Hence the country with the higher rate of interest is expected to have the depreciating currency. The forward premium anomaly found in many studies is that a regression of the form of,

$$s_{t+k} - s_t = á + \beta(i_t^* - i_t) + u_{t+k},$$

is found to have a negative slope coefficient, β, which implies that the country with the higher rate of interest is expected to have a currency appreciation.

A generalization of equation (2) is to specify real returns over the current and future consumption stream, so that equation (3) is modified to the discrete time asset pricing approach of Lucas (1978) to be,

$$(4) \qquad E_t[(F_t - S_{t+k})/P_{t+k}]U^{\prime}(C_{t+k})/U^{\prime}(C_t) = 0,$$

where $U^{\prime}(C_{t+1})/U^{\prime}(C_t)$ is the marginal rate of substitution in terms of utility derived from current and future consumption. On assuming a logarithmic utility function with a constant coefficient of relative risk aversion, (CRRA), denoted by γ, then equation (4) can be expressed as,

$$(5) \quad E_t s_{t+1} - f_t = -(1/2)Var_t(s_{t+1}) + Cov_t(s_{t+1}p_{t+1}) + \gamma Cov_t(s_{t+1}c_{t+1}).$$

The last term, will be denoted by $\rho_t = \gamma Cov_t(s_{t+1}c_{t+1})$ and is known as a time dependent risk premium. There are several possible theoretical developments of this term. For example, if each country's consumption growth is assumed to be proportional to world income growth so that $C_t = Y_t^{\gamma}$, then under the CRRA assumption, the risk premium term in (6) is $\rho_t = Cov_t(s_{t+1}y_{t+1}^{\gamma})$. More generally, the risk premium is frequently expressed as, $\rho_t = \rho Cov_t(s_{t+1}q_{t+1})$, where q_t is the logarithm of the intertemporal marginal rate of substitution.

It should be noted that in many previous studies of time varying risk premium, e.g. Domowitz and Hakkio (1985), Hodrick (1987, 1989) and Kaminsky and Peruga (1990), the x_t variables typically contain conditional variances and covariances of the asset price vis-à-vis the forcing variables. However, the fact that goods market variables needed for the Lucas-Breeden model such as relative prices, consumption levels, money supplies, etc are generally only observed monthly and the fact that goods market prices are very smooth compared to asset market prices makes estimation and testing these relationships problematic. Also, as noted by Baillie and Bollerslev (1989), exchange rates typically possess little ARCH effects at the monthly level, so that studies such as Kaminsky and Peruga (1990) have been unsuccessful in collaborating the Lucas Breeden asset pricing model. Consequently, it is attractive to estimate the model from daily data where the effects of changes in the intervention variables may be more clearly apparent.

Data

This study uses data provided by the Board of Governors of the Federal Reserve System and the German Bundesbank from January 3, 1987 through January 22, 1993. This includes the periods around the Louvre Accord and the stock market crash in October, 1987. For each country, the data record in 100 million US dollar units the actual net purchases of US dollars vis a vis the German mark from close of business on day t-2 to close of business on day t-1.

The exchange rates are recorded at 9:30am Paris time as supplied to us by Olsen & Associates of Zurich, Switzerland. For each bilateral exchange rate there are four intervention variables: the purchases and sales of dollars by each country. The intervention variables are aligned so as to be predetermined with respect to the change in the exchange rate from day t to t+1.

We use a unique data set on overnight Eurocurrency deposit rates obtained from the Paris market through DRIFACS, with the ultimate source being Credit Lyonnais, Paris. The use of this data, which are essentially interest rates of one day maturity time, allows us to avoid many of the econometric problems associated with forward rates with overlapping contracts, which reduce the efficiency in tests of unbiased expectations of the forward rate. With one month forward contracts it is important to elaborately match in accordance with the settlement conventions as described by Riehl and Rogriguez (1977), Levine (1989) and Bekaert and Hodrick (1993).

Figure 1 shows the movement in the DM/$ rate and the overnight interest rates during the sample period. The U.S. rate exceeded that of Germany until Fall 1990. UIRP requires that higher(lower) U.S. rates be offset by an expected decline(rise) in the DM/$. However, causality could run in either direction. The DM/$ began a steady decline from over 2 in June 1989. Figure 2 shows the logarithmic analogues which together comprise the deviation from UIRP. Clearly, and not surprisingly, the volatility in the exchange rate exceeds that of the interest rates. Figure 3 shows the logarithmic change in the exchange rate together with U.S. intervention vis-à-vis the DM, with positive intervention indicating purchases of dollars.

It is not clear from Figure 3 whether U.S. intervention was intended to affect either the level or the volatility of the DM/$. There were only two possibly distinct intervals of intervention activity-the period around the Louvre Accord of February 20-21, 1987 and the equity market crash of October 1987.

An Econometric Model for UIRP

An econometric model for the ex post deviations from uncovered interest rate parity is given by,

(6) $y_t = (\Delta s_{t+1} - i_t^* + i_t) = b' x_t + \varepsilon_{t+1},$

(7) $\varepsilon_t = \xi_t \sigma_t,$

(8) $\sigma^2_t = \omega + \beta\sigma^2_{t-1} + (1 - \beta L - (1 - \varphi L)(1 - L)^{\delta})\varepsilon^2_t,$

where ξ_t is i.i.d.(0,1) process, b are a k dimensional vector of predetermined variables, and σ^2_t is a time-varying, positive and measurable function of the information set at time t-1. Hence the conditional variance σ^2_t, is represented by a FIGARCH (Fractionally Integrated AutoRegressive Conditional Heteroskedastic) models as developed by Baillie, Bollerslev and Mikkelsen (1996). The above model is the FIGARCH(1,δ,1) process and generates the type of very slow decay which are frequently observed in the autocorrelations of squared returns, absolute returns and the power transformations of returns; see Ding, Granger and Engle (1993) and Ding and Granger (1996). The general FIGARCH(p,δ,q) process is given by,

(9) $$[1 - \beta(L)]\sigma^2_t = \omega + [1 - \beta(L) - \varphi(L)(1 - L)^{\delta}]y^2_t,$$

where $\beta(L) = \beta_1 L + ... + \beta_p L^p$, $\varphi(L) = 1 - \alpha(L) - \beta(L)$, and $\alpha(L) =$ $á_1 L + ...+ á_q L^q$; while ä denotes the long memory, or fractional parameter and is defined for $0 < \delta < 1$. There are some further important restrictions on the parameters; namely that $\omega > 0$ and that all the roots of $[1 - \beta(L)]$ and $\varphi(L)$ must lie outside the unit circle. By straightforward algebra, the process can also be expanded as the infinite order ARCH model,

(10) $$\sigma^2_t = \omega/[1 - \beta(1)] + \lambda(L)y^2_t,$$

where

(11) $$\lambda(L) = \{1 - [1 - \beta(L)]^{-1}\varphi(L)(1 - L)^{\delta}\}.$$

The key property of the above FIGARCH model, which distinguishes it from alternatives is that it again implies very slow hyperbolic rate of decay on the impulse response weights $\lambda_k \approx k^{d-1}$ in equations (10) and (11), which is essentially the "long memory" property, or "Hurst effect". Many well known ARCH models are special cases of the FIGARCH representation. For example, when d = 0, p = q = 1, then equation (8) reduces to the GARCH(1,1) model. When d = p = q

= 1, in equation (8) realizes the Integrated GARCH, or IGARCH(1,1) model, and implies complete persistence of the conditional variance to a shock in squared returns. The attraction of the flexibility of the FIGARCH process is that intermediate ranges of persistence can be introduced by having d in the range, $0 < d < 1$. In many practical situations quite low order models are adequate, such as the FIGARCH(1,d,1) model. An even simpler model to be applied later is the FIGARCH(1,d,0) process,

$$(12) \qquad (1 - \beta L)\sigma^2_t = \omega + [1 - \beta L - (1 - L)^d]y^2_t.$$

For this model the impulse response weights in (6) are $\sigma^2_t = \omega/(1 - \beta) + \lambda(L)y^2_t$, and it can be shown that $\lambda_k = \Gamma(k+d-1)/\{\Gamma(k)\Gamma(d)\}[(1-\beta)-(1-d)/k]$, where $\Gamma(.)$ is the gamma function. For large lag k, $\lambda_k = [(1-\beta)/\Gamma(d)]k^{d-1}$, which generates the same slow hyperbolic rate of decay on the impulse response weights of the conditional variance σ^2_t.

In this study the estimation of the system of equations in (6) through (8) is facilitated by means of Quasi Maximum Likelihood Estimation (QMLE); see Bollerslev and Wooldridge (1988). The procedure uses conventional non-linear procedures to maximize the Gaussian log likelihood function,

$$(13) \ \log(\theta) = -(T/2)\log(2\Pi) - (1/2)\Sigma_{t=1,T}[\log(\sigma^2_t + \varepsilon^2_t\sigma^{-2}_t],$$

with respect to a specified vector of parameters, θ. It should be noted that the numerical procedures are quite general and can be readily extended to models such as the regression model with ARMA disturbances and FIGARCH volatility process. Since most return series are not well described by the conditional normal density in (14); subsequent inference using robust standard errors, is based upon noting that the limiting distribution of the QMLE are given by,

$$(15) \qquad T^{1/2}(\theta_T - \theta_0) \Rightarrow N\{0,\ A(\theta_0)^{-1}B(\theta_0)A(\theta_0)^{-1}\},$$

where A(.) and B(.) represent the Hessian and outer product gradient respectively; and θ_0 denotes the true parameter values.

Results

Some tests of uncovered interest rate parity are obtained by regressing the spot rate returns on the interest rate differential and also the same menu of predetermined variables as in equation (6). Table 1 presents the results estimating the model when four intervention variables (buying and selling of dollars vis-à-vis the DM by both countries) are in the conditional mean and when a GARCH(1,1) formulation is employed for the conditional variance. Results for equation (6) are the same with a FIGARCH specification or when day-of-the-week dummies are included in the conditional mean as is discussed later. In all cases the estimate of the coefficient associated with the interest rate differential was negative in accord with the average value of -0.88 found in 75 separate studies by Froot and Thaler (1990). However, with the overnight Euro deposit rate data, the .95 percentile confidence intervals around the estimated value of b were sufficient to include the value of b = 1. Since the main focus in this study is that of the risk premium the value b = 1 is maintained throughout so that the dependent variable is the ex post return over uncovered interest rate parity, $y_t = (\Delta s_{t+1} - i_t^* + i_t)$. Since the sampling interval of one day exactly matches the maturity time of the forward contract, the y_t series appears uncorrelated.

Consequently, table 2 reports estimates of the following model,

$$(6') \quad [s_{t+1}-(i_t^* - i_t)] = \mu + \gamma_1 US_t^b + \gamma_2 US_t^s + \gamma_3 G_t^b + \gamma_4 G_t^s + \Sigma_{j=1,4}\lambda_j D_{jt} + \varepsilon_t,$$

(7) $\varepsilon_t = \xi_t\sigma_t$,

(8) $\sigma^2_t = \omega + \beta\sigma^2_{t-1} + (1 - \beta L - (1 - \varphi L)(1 - L)^{\delta})\varepsilon^2_t$,

where ξ_t is the i.i.d.(0,1) process, σ^2_t is the conditional variance process, D_{jt} denote day of the week dummy variables. Baillie and Bollersev (1989), Hsieh (1989), and McFarland, et al. (1982) have discussed the possibility of day-of-the-week effects in the conditional mean and conditional volatility of daily exchange rate returns.

US^b_t denotes the Federal Reserve Bank buying dollars, US^s_t denotes the Federal Reserve Bank selling dollars, and G^b_t and G^s_t denotes corresponding actions by the German Bundesbank. The introduction of intervention in this manner is consistent with viewing intervention as providing a signal of policy. U.S. purchases of dollars, by this reasoning, should signal that the authorities have information, presumably about policy, which once known, would boost the DM/$. This applies to the impact of Bundesbank dollar purchase as well. Thus we would expect estimates of both γ_1 and γ_3 to be positive and for estimated γ_2 and γ_4 to be negative.

Table presents results from estimating the system (6'), (7), and (8). Diagnostic tests indicate the success of the FIGARCH(1,d,1) specification in modeling the conditional variance of the deviations from UIRP and in offering a significant improvement over alternatives such as GARCH. None of the daily dummies (Monday through Thursday) were significant in the conditional mean. Of the four intervention variables, only German buying has a significant impact with the opposite sign from that implied by the signaling hypothesis. This is consistent with German dollar purchases reducing DM/$.

The second column of Table 2 indicates the results for a subperiod that spans just after the October 1987 crash through the end of 1989. The coefficient on German purchases continues to be negative but U.S. buying has a positive impact.

Conclusion

This paper has presented preliminary results of an investigation into the impact of U.S. and German dollar intervention vis-à-vis the DM on the deviation from uncovered interest rate parity (UIRP). We view intervention as possibly signaling future policy so that dollar purchases should increase the deviation from UIRP. The approach adopted differs from previous work by utilizing overnight Eurocurrency rate data matched exactly to exchange rates. A FIGARCH formulation of the conditional variance and thus the standard errors is implemented with QMLE.

We find that German dollar purchases decreased the deviation from UIRP, a result consistent with our previous finding (Baillie and Osterberg [1997a]) that it decreased the DM/$ during August 6, 1985 through March 1, 1990. This result holds up in an examination of October 20, 1987 through 1989 subperiod during which, however, we find a positive impact of U.S. dollar purchases. This latter result is consistent with our analysis of the risk in the forward market (Baillie and Osterberg[1997b]). The same conclusion results when we account for possible differences between coordinated and unilateral interventions.

In general, these results are negative for the signalling hypothesis. However, additional results have confirmed similar findings for analyzing the impact of cumulative intervention. Further work might take account of market conditions and policy intentions during specific subperiods. For example, intervention at times might be intended to ‘lean against the wind’, or to reduce volatility.

Table 1: Estimation of the model:

$\Delta s_{t+1} = \mu + b(i_t^* - i_t) + \gamma_1 US_t^b + \gamma_2 US_t^s + \gamma_3 G_t^b + \gamma_4 G_t^s + \varepsilon_t,$

$\varepsilon_t = \xi_t \sigma_t,$

$\sigma^2_t = \omega + \alpha\varepsilon^2_{t-1} + \beta\sigma^2_{t-1}$

Parameters	
μ	-0.256(0.203)
b	-1.813(2.210)
γ_1	0.009(0.007)
γ_2	-0.003(0.007)
γ_3	-0.011(0.005)
γ_4	0.005(0.005)
ω	1.728(0.626)
α	0.081(0.015)
β	0.893(0.019)
m_3	0.232
m_4	4.430
Q_{20}	14.88
Q^2_{20}	19.14
ln(∴)	-5029.86

Key: There are T = 1,463 observations from January 5, 1987 through January 22, 1993.

Table 2: Estimation of the model:

$[\Delta s_{t+1}-(i_t^*-i_t)] = \mu + \gamma_1 US_t^b + \gamma_2 US_t^s + \gamma_3 G_t^b + \gamma_4 G_t^s + \Sigma_{j=1,4}\lambda_j D_{jt} + \varepsilon_t,$

$\varepsilon_t = \xi_t \sigma_t,$

$\sigma^2_t = \omega + \beta\sigma^2_{t-1} + [1 - \beta L - (1 - \varphi L)1 - L)^{\delta}]\varepsilon^2_t,$

Parameters	Full sample	Oct 20,1987 -Dec 31, 1989.
μ	-0.408(0.391)	0.640(0.690)
γ_1	0.009(0.007)	0.022(0.009)
γ_2	-0.004(0.007)	-0.010(0.005)
γ_3	-.010(.005)	-0.014(0.005)
γ_4	0.005(0.005)	0.007(0.004)
λ_1	0.382(0.681)	-1.826(.926)
λ_2	0.080(0.558)	-0.935(0.904)
λ_3	0.287(0.545)	0.863(0.885)
λ_4	0.323(0.526)	-1.647(0.979)
δ	0.525(0.144)	0.624(0.290)
ω	1.682(0.804)	1.815(1.544)
β	0.671(0.103)	0.713(0.176)
φ	0.182(0.069)	0.042(0.140)
m_3	0.17	0.12
m_4	4.21	3.74
Q_{20}	15.43	15.16
Q_{20}^2	19.37	14.93
ln(∴)	-5028.89	-1783.370

Key: There are T = 1,463 observations for January 5, 1987 through January 22, 1993. In the reduced sample, T = 530. Asymptotic robust standard errors are in parentheses. The Q_m statistic is the Ljung-Box test for autocorrelation based on the first m autocorrelations of the standardized residuals. Q_m^2 is the

Ljung-Box test for ARCH effects based on the first m lags of the autocorrelations of the squared standardized residuals.

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FIGURE 1: U.S. AND GERMAN OVERNIGHT RATES AND DM/$ EXCHANGE RATE

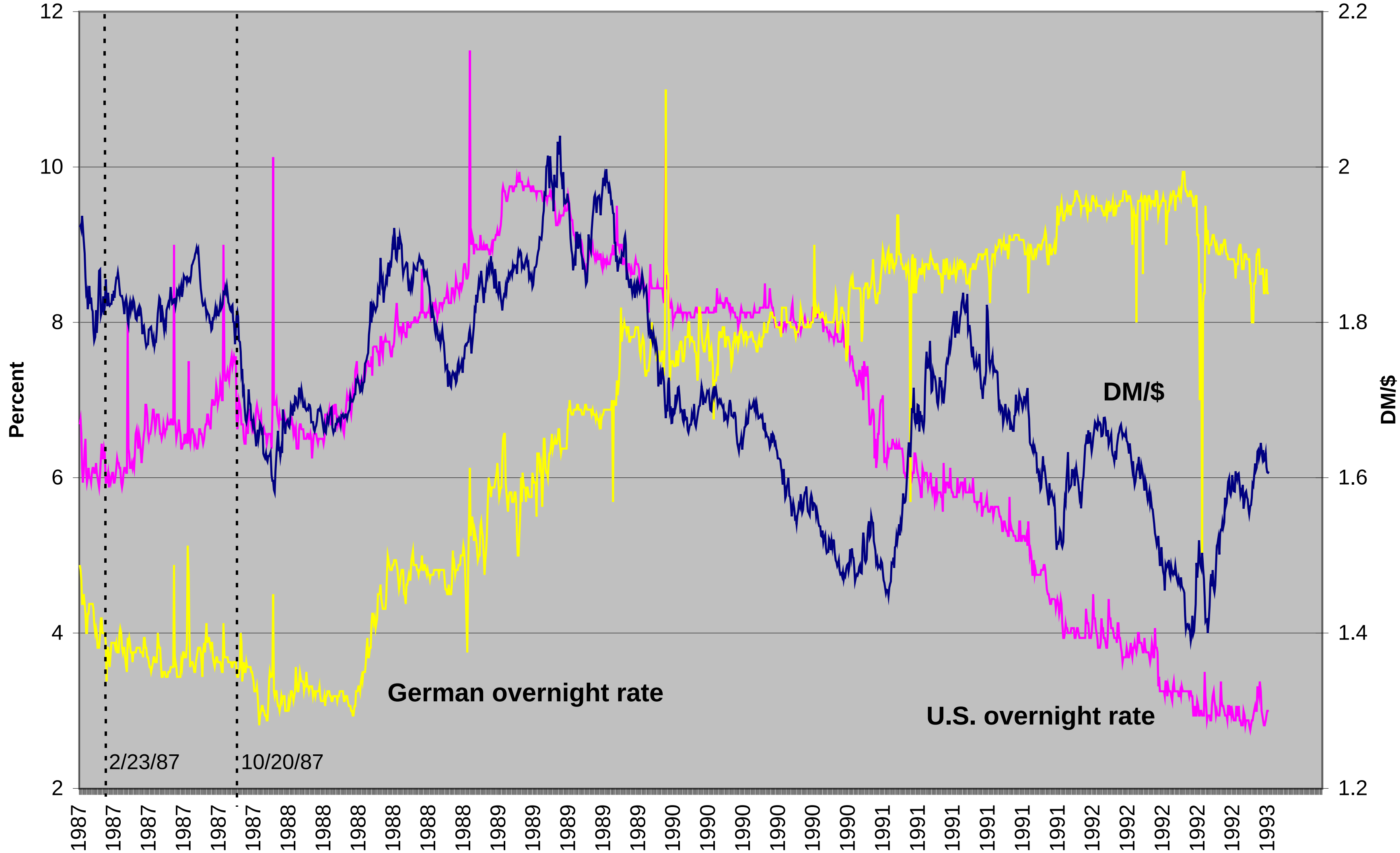

FIGURE 2: 1000(Δln dm/$), 1000(ln(1+German rate/36000)-ln(1+U.S. rate/36000))

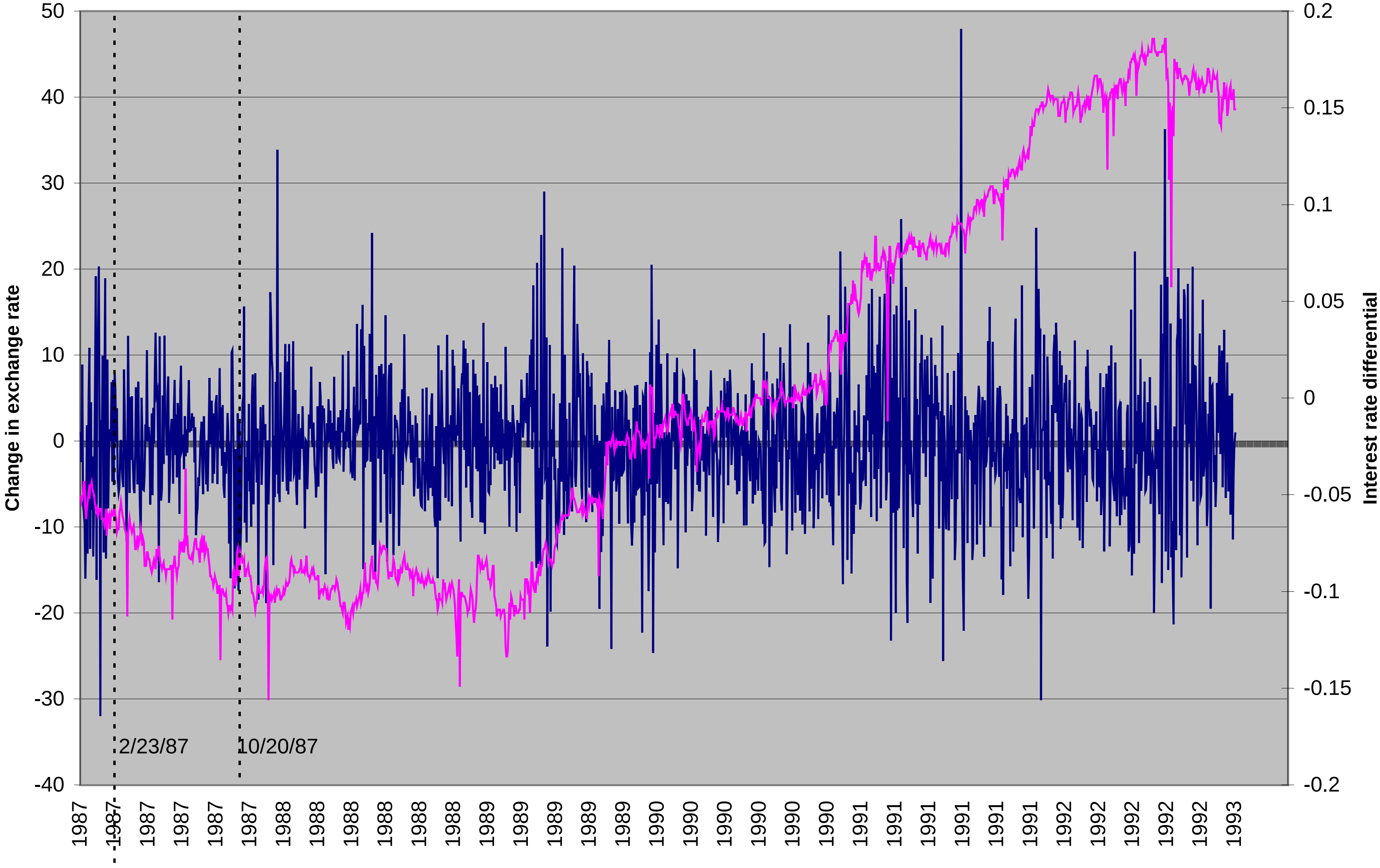

FIGURE 3: 1000(Δln dm/$), U.S. Intervention vis-à-vis dm