

Japanese Foreign Exchange Intervention and the Yen-to-Dollar Exchange Rate: A Simultaneous Equations Approach Using Realized Volatility

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Abstract

We use realized volatility to study the influence of Japanese central bank interventions on the yen-to-dollar exchange rate. A system of equations for returns, logarithmic realized volatility, and interventions provides a comprehensive view on the problem without endogeneity bias, unlike earlier latent variable specifications. We find that during the period 1991 through 1995, interventions of the Japanese monetary authorities could not move the yen-dollar rate into the desired direction. We measure an increase in volatility associated with interventions. During the period 1995 through 1998, the estimations are consistent with interventions that successfully influenced returns. After 1998 up to the last intervention episode in 2004, interventions did not have a significant impact on returns but reduced realized exchange rate volatility.

Key Words: Realized Volatility, Structural Change, GMM, Foreign Exchange Intervention, Japan

JEL Classification: C32, E58, F31, F33, G15

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1 Introduction

Since the Japanese monetary authorities have released data on their foreign exchange intervention activities in 2001, a steadily increasing number of studies have scrutinized the effects of Japanese foreign exchange intervention. One of the main challenges to address is an endogeneity problem: If there is significant correlation between interventions and exchange rate returns or volatility, does this support the hypothesis that interventions cause changes in exchange rate movements or does this support the reverse direction that exchange rate movements trigger interventions? Building upon the paper of Dominguez (1998), the studies of Ito (2003), Frenkel, Pierdzioch and Stadtmann (2005), Watanabe and Harada (2006), and Hillebrand and Schnabl (2006), among others, have used daily time series in Generalized Autoregressive Heteroskedasticity (GARCH) frameworks to study the impact of Japanese foreign exchange intervention on the volatility of the yen-to-dollar exchange rate. Instead of trying to measure the success of interventions in pushing the exchange rate to a desired level, these studies use the smoothing of exchange rate volatility as a success criterion. The coefficients in the GARCH mean equations were rendered not interpretable because of endogeneity bias. Separate estimations of reaction functions for the monetary authorities usually found that interventions correlated with exchange rate returns but not with volatility. Therefore, endogeneity did not seem to be a problem in the conditional volatility equation, at least not if one accounted for linear effects only. These studies found mixed evidence that Japanese foreign exchange interventions have increased or decreased exchange rate volatility, depending on the time period.

As the GARCH time series approaches have not been able to fully resolve the endogeneity issue, a new strand of literature has evolved that has used event studies to analyze the success of Japanese foreign exchange intervention (Neely 2005). Fatum and Hutchison (2003) separate intervention episodes and analyze the subsequent effects on the exchange rate. They find evidence in favor of successful Japanese intervention, as mean exchange rate changes after

interventions are significantly smaller than mean pre-intervention changes.

Kearns and Rigobon (2005) specify a multiple equation model for returns and interventions that uses a change in intervention policies to identify the parameters. Kim (2003) proposes a structural VAR model for returns and interventions and estimates the effects of intervention and monetary policy with monthly data.

The concept of realized volatility, introduced by Andersen and Bollerslev (1998) and a very active area of volatility research (Andersen, Bollerslev, Diebold, and Labys 2001, 2003, Andersen, Bollerslev, Diebold, and Ebens 2001, Barndorff-Nielsen and Shephard 2002, Zhang, Mykland and Ait-Sahalia 2005, Barndorff-Nielsen, Hansen, Lunde, and Shephard 2007, among others), allows us to consider volatility as an observed rather than as a latent variable, albeit with measurement error. Therefore, we can specify a system of equations that contains not only returns and interventions but also realized volatility. Such a system provides a comprehensive framework to study the interplay of the first and the second moment of the return distribution of the yen-to-dollar rate with interventions without endogeneity bias. Realized volatility has been used in the intervention literature to study the jump component in foreign exchange volatility, see Beine, Lahaye, Laurent, Neely, and Palm (2007).

We break our total sample period from 1991 through 2004 up into three segments corresponding to changes in the intervention strategy of the Japanese authorities from frequent smaller interventions to infrequent larger ones in 1995 (Itô and Yabu 2004) and corresponding to a broad liberalization of the foreign exchange market in Japan in 1998 (Itô and Melvin 1999). Generalized Method of Moments (GMM) estimation of the system of equations on the sub-periods reveals that Japanese foreign exchange interventions were not successful during the period 1991 through 1995, neither in influencing the returns nor in reducing the volatility of the yen-to-dollar rate. On the contrary, we measure a significantly positive coefficient of interventions in the volatility equation. During the period 1995 through 1998, the estimations are consistent with interventions that

successfully influence exchange rate returns. During the period 1998 through 2004, interventions did not influence returns but significantly reduced volatility. The results therefore indicate a change toward a more successful intervention policy.

We study the robustness of our findings using unrestricted vector autoregressions as alternative model and estimation frameworks as well as alternative structural specifications. We also use a statistical change-point detector on the realized volatility time series to study the sensitivity to different change-points. We find that most of our results, in particular the reduction of exchange rate volatility during more recent intervention periods, are robust to different approaches.

The outline of the paper is as follows. In Section 2, we describe our data set. In Section 3, we describe the model specification, the estimation method, the choice of change-points, and the main findings. In Section 4, we study the robustness of our results to alternative vector autoregressive model specifications, generalizations of our model that allow for other financial variables, and alternative change-point settings. Section 5 concludes.

2 Data

The main time series considered in this paper are 5-minute intra-day quotes of the yen-to-dollar exchange rate and the daily time series of interventions reported by the Japanese authorities. The sample period is 2-Apr-1991 through 27-Oct-2004. This corresponds to a sample size of 3530 days. The choice of the sample size is motivated by the availability of intervention data from April 1991 and the latest interventions, which at the time of writing this paper took place in 2004.

High-frequency exchange rate data are purchased from Olsen Financial Technologies. The data are filtered for outliers and 5-minute prices are generated by linearly interpolating the average of log-bid and log-ask for two closest ticks.

We delete the weekends from Friday 21:05 Greenwich Mean Time (GMT) until Sunday 21:05 GMT, holidays, and days of unusually low trading volume.

Let $X(t)$ denote the log of the yen-to-dollar exchange rate at time t . We compute the *realized variance* time series σ_t^2 from the logarithmic 5-minute returns of the yen-to-dollar exchange rate by the sample version of the quadratic variation $\langle X \rangle(t) - \langle X \rangle(t-1)$ over one day:

$$\sigma_t^2 := \sum_{j=1}^n |X(\tau_j) - X(\tau_{j-1})|^2, \quad (1)$$

on the partition $\Pi = \{\tau_0 = t-1, \tau_1, \dots, \tau_n = t\}$ of the interval $[t-1, t]$ with mesh $|\Pi| = 5$ minutes.

For a semi-martingale $Y(t)$, as $|\Pi| \rightarrow 0$, it follows from the Itô isometry that the quadratic variation is given by

$$\langle Y \rangle(t) - \langle Y \rangle(t-1) = \lim_{|\Pi| \rightarrow 0} \sum_{j=1}^n |Y(\tau_j) - Y(\tau_{j-1})|^2 = \int_{t-1}^t \sigma_Y^2(s) ds,$$

which is *integrated variance* and the object of interest. Here, $\sigma_Y^2(t)$ is the instantaneous variance of Y . Barndorff-Nielsen and Shephard (2002) show for prices of financial assets that market microstructure noise makes realized variance a noise-contaminated estimator of integrated variance. Several consistent estimators of integrated variance that use tick-level data have been proposed (Zhang, Mykland, and Ait-Sahalia 2005, Hansen, Large, and Lunde 2007, Barndorff-Nielsen, Hansen, Lunde, and Shephard 2007). We use the simple sparse sampling estimator since our data set is aggregated to 5-minute quotes (Andersen and Bollerslev 1998, Andersen, Bollerslev, Diebold, and Labys 2001). Realized volatility σ_t is obtained by taking the square root of daily realized variance.

The daily interventions of the Japanese monetary authorities are reported on the web site of the Japanese Ministry of Finance. The exact intervention time, the number of interventions within a day, the intervention market (Tokyo, London, or New York), and the exchange rate at the time of intervention remain undisclosed. The reported amounts are in billion yen; we convert them into

billion dollars based on daily exchange rates. The interventions are stamped at Tokyo time. The high-frequency quotes of the yen-to-dollar exchange rate are stamped at Greenwich Mean Time, which is Tokyo time minus nine hours. Therefore, interventions I_t at Tokyo time precede the returns r_t and the realized volatility σ_t^2 at Greenwich Mean Time. Sometimes the Federal Reserve intervenes on behalf of the Japanese authorities. There are no publicly available data on these transactions.

We also consider U.S. foreign exchange interventions in the yen-to-dollar market; the data are provided by the Federal Reserve Board. We include them into our comprehensive intervention time series since the U.S. interventions were almost without exception coordinated with the Japanese authorities. That is, they occurred at the same day and with the same sign. The U.S. interventions make up only a very small fraction in this sample. Table 1 reports summary statistics for the interventions of the Bank of Japan and the Federal Reserve. Figures 1, 2, and 3 show the time series of pooled Japanese and Federal Reserve interventions, the daily returns on the yen-to-dollar exchange rate, and the logarithmic daily realized volatility of the yen-to-dollar exchange rate.

Other time series used for robustness checks are daily index prices of the Nikkei 300 (Bloomberg series NEY), the Federal Funds Rate (Bloomberg: FDFD), and the Japanese uncollateralized overnight interbank interest rate (Bloomberg: JYMU1T). The Japanese overnight rate is available only after 11-Apr-1996.

FIGURES 1, 2, 3 ABOUT HERE

3 A Simultaneous Equations Model of Exchange Rate Moments and Intervention

In this section, we specify and estimate the system $y_t := (r_t, \log \sigma_t, I_t)$, where r_t are the daily log returns of the yen-to-dollar exchange rate, σ_t is the daily

realized volatility of the yen-to-dollar exchange rate, and I_t are combined interventions by Japanese and U.S. monetary authorities.

3.1 Specification

We consider the following system of equations

$$r_t = \alpha_0 + \alpha_1 I_t + u_t, \quad (2)$$

$$\log \sigma_t = \beta_0 + \beta_1 \log \sigma_{t-1} + \frac{\beta_2}{5} \sum_{j=1}^5 \log \sigma_{t-j} + \frac{\beta_3}{21} \sum_{j=1}^{21} \log \sigma_{t-j} + \beta_4 |I_t| + v_t, \quad (3)$$

$$I_t = \gamma_1 I_{t-1} + \gamma_2 r_t + \gamma_3 r_{t-1} + w_t, \quad (4)$$

where $r_t = X_t - X_{t-1}$ are the daily log returns of the yen-to-dollar exchange rate, σ_t is the daily realized volatility as defined in equation (1), and I_t are the interventions ($I_t > 0$: dollar purchases, $I_t < 0$: dollar sales). We will use GMM to estimate the system, so the only assumption we need on the error processes u , v , and w is that they are such that the system y is stationary and ergodic.

The first equation models the influence of interventions on yen-to-dollar returns. Purchases of dollar (sales of yen) by the Japanese authorities, intended to depreciate the yen against the dollar, should increase returns r_t , meaning that one dollar buys more yen. The opposite should hold for sales of dollars (purchases of yen).

The second equation specifies the log of realized volatility as a function of its own lags in the spirit of the Heterogeneous Autoregressive Realized Volatility (HAR-RV) model (Corsi 2004, Andersen, Bollerslev, and Diebold 2007). This structure can capture the stylized fact of long memory that is found in realized volatility. The specification is chosen such that the standard stationarity condition $\beta_1 + \beta_2 + \beta_3 < 1$ for autoregressive models holds. Interventions enter in absolute value. Note that while realized volatility and the second moment of r_t are related, we do not impose any condition that the second variable is the variance of the first variable in any sense. The reason for not doing this is that realized variance is an estimator for integrated variance, which in turn

is an estimator for conditional variance (see Andersen, Bollerslev, Diebold, and Labys 2001, equation (11a)). Relating log realized volatility in (3) to conditional variance in (2) would imply a nonlinear relation between the two equations. It would add an error term due to the estimator character of realized variance for conditional variance. Omitting this relationship will lead to nonlinear dependence of u_t and v_t :

Lemma 1. *Let z_t be the set of instruments in equation (2) and let $u_t = \sqrt{\text{Var}(r_t|I_t)}\varepsilon_t$, where $\varepsilon_t \sim WN(0,1)$ and $\mathbb{E}(z_t\varepsilon_t) = 0$. Let realized variance be an unbiased estimator of $\text{Var}(r_t|I_t)$, i.e.,*

$$\sigma_t^2 = \text{Var}(r_t|I_t) + \eta_t,$$

where η_t is a zero-mean random variable uncorrelated with ε_t and z_t . Let realized variance be described by $\log \sigma_t = f(\{\sigma_\tau\}_{\tau \leq t-1}, I_t) + v_t$, where f is some function measurable in its arguments (for example as in equation (3)) and v_t is a zero-mean process uncorrelated with ε_t and z_t . Then,

$$\mathbb{E}(z_t u_t) = 0.$$

Proof. Write u_t as

$$u_t = \sqrt{\sigma_t^2 - \eta_t} \varepsilon_t,$$

using realized variance as estimator for conditional volatility. Then,

$$\begin{aligned} \mathbb{E}(z_t u_t) &= \mathbb{E}\left(\sqrt{\sigma_t^2 - \eta_t} z_t \varepsilon_t\right), \\ &= \mathbb{E}\left(\sqrt{\exp(2f(\{\sigma_\tau\}_{\tau \leq t-1}, I_t) + 2v_t) - \eta_t} z_t \varepsilon_t\right), \\ &= \mathbb{E}\sqrt{\exp(2f(\{\sigma_\tau\}_{\tau \leq t-1}, I_t) + 2v_t) - \eta_t} \mathbb{E}(z_t \varepsilon_t) = 0. \end{aligned}$$

□

Lemma 1 shows that GMM estimation is suited to handle the omitted nonlinear dependence of u and v . The situation is as described in Hall (2005), p. 45: the GMM estimator remains consistent and asymptotically normal.

The third equation is the reaction function of the intervening authorities. Interventions exhibit strong clustering, therefore a first-order autoregressive term is included. We have included two lags of returns on the yen-to-dollar exchange rate as determinants of interventions. It has been documented in many studies that Japanese authorities do not target exchange rate volatility in intervention decisions, therefore we do not include volatility in the third equation (Dominguez 1998, Itô 2003, Frenkel, Pierdzioch, and Stadtmann 2005, among others).

Our goal in modeling the system $y_t = (r_t, \log \sigma_t, I_t)$ in equations (2) to (4) is to choose the simplest and most parsimonious specification that captures the effects of intervention. We will show in Section 4 that none of the assumptions that we make here on the determinants of the three variables nor the estimation method that we choose in Section 3.2 drive our main results. The main findings of this study can also be obtained in an unrestricted vector autoregression.

Before the theory of realized volatility was available, equations (2) and (3) were usually specified in a GARCH framework with interventions as exogenous variables. Equation (4), the reaction function of the monetary authorities, had to be estimated separately. Examples for studies that follow this approach are Dominguez (1998), Bonser-Neal and Tanner (1996), Frenkel, Pierdzioch, and Stadtmann 2005, and Hillebrand and Schnabl (2006), among others. In this setup, volatility was latent and the equations (2) and (3) of the GARCH regression suffered from simultaneous equation bias because equation (4) was not part of the system. Separate estimations of the reaction function (4) routinely indicated that interventions were triggered by changes in returns, underlining the endogeneity problem in equation (2). The conditional volatility equation on the other hand seemed to be statistically fine since volatility (squared daily returns or fitted GARCH series) did not seem to influence interventions in the reaction function estimation. Therefore, the estimated coefficients of the mean equation of the GARCH model could not be interpreted and measuring success of interventions was limited to measuring changes of volatility in response to

interventions.

Realized volatility allows us to treat σ_t as an observed variable rather than as a latent variable and set up a system of equations. Multiple equation models have been employed before to analyze the effects of interventions on exchange rates (Kim 2003, Kearns and Rigobon 2005, Neely 2005). Our contribution to this literature is that we include volatility in the system and therefore disentangle the interplay of returns, volatility, and interventions. This resolves the endogeneity problem of the approach using GARCH and exogenous interventions since the reaction function can be estimated simultaneously with the return and volatility function.

3.2 Estimation Method

The parameter vector to be estimated is

$$\theta = (\alpha_0, \alpha_1, \beta_0, \beta_1, \beta_2, \beta_3, \beta_4, \gamma_1, \gamma_2, \gamma_3).$$

We cannot make standard distribution assumptions on the error terms because the intervention time series is equal to zero most of the time and has pronounced clusters of large interventions (Figure 1). We therefore estimate the system by GMM, which does not require a specific error distribution to derive inferences (Hansen 1982, Hall 2005).

We choose interventions at lags one and two as instruments for the first equation. Many studies have shown that daily intervention data have significant low order autocorrelations and the first few lags are routinely included in the specification of reaction functions (e.g., Ito 2003, Dominguez 1998). The sample partial autocorrelation function for the Japanese intervention series drops off after the first two lags. Therefore, the assumption that the instruments are orthogonal to the error in the regression of interventions I_t on the instruments seems plausible.

For the second equation, we choose the absolute value of interventions at lags one and two as instruments for the absolute value of concurrent interventions.

The motivation is the same as for the first equation. We also add the lags of log realized volatility according to the HAR-RV-type specification to the set of instruments for this equation.

For the third equation, we choose interventions at lags one and two as well as log realized volatility at lags one and two as instruments. The latter identify the return lags that are included in the equation since the returns correlate with realized volatility (about -10% contemporaneous correlation) and authorities usually do not monitor volatility as parameter for interventions.

We use a heteroskedasticity and autocorrelation consistent estimator with Bartlett spectral kernel for the covariance matrix of the moment conditions and bandwidth selected according to Newey and West (1994). Instrument weakness can be an issue in financial data (Stock, Wright, and Yogo 2002). An advantage of the system as we have specified it so far is that it is completely self-contained; no variables outside the system y are used. We relax this in Section 4 and include other financial variables listed in the Data Section 2 to check robustness of our results. We also consider unrestricted vector autoregressions as well as other possible channels of intervention such as interest rates. We show that the main findings of this study do not depend on the specific equation structure (2) to (4) and the choice of instruments.

3.3 Structural Breaks

The periods of high and low volatility that can be seen in Figure 3 can be understood as different parameter regimes interrupted by structural breaks, that is, changes in the data-generating parameters of the volatility model under study. The possibility of structural breaks and its implications for the estimation of serial correlation and persistence in volatility has been discussed widely (Andreu and Ghysels 2002, Bos, Franses, and Ooms 1999, Diebold and Inoue 2001, Granger and Hyung 2004, Hillebrand 2005, Lamoureux and Lastrapes 1990, LeBaron 2001, Rapach and Strauss 2008).

Earlier studies in the intervention literature have found evidence for structural breaks in the yen-dollar exchange rate. In this study, we will apply two exogenously specified breaks: The first is 1-Jun-1995 and the second is 1-Apr-1998. The break in June of 1995 is identified in Itô and Yabu (2004). At this point, Japanese authorities changed their approach to intervention from frequent small-size interventions to infrequent large-size interventions. Itô and Yabu (2004) show that the reaction function of the Japanese authorities changed at that time. The break in April 1998 is identified in Itô and Melvin (1999). It marks a broad deregulation of the foreign exchange market in Japan. Itô and Melvin (1999) estimate that conditional volatility changed at this point in time. In Section 4, we apply a statistical change-point detector to the time series of realized volatility to check our findings for robustness to different choices of change-points.

3.4 Main Findings

Tables 2 through 4 report the estimated coefficients of model (2) through (4) on the three resulting segments of data: (1) 2-Apr-1991 to 1-Jun-1995, (2) 2-Jun-1995 to 1-Apr-1998, and (3) 2-Apr-1998 to 27-Oct-2004.

TABLES 2, 3, 4 ABOUT HERE

There are two concepts of “success” of interventions discussed in the literature: Either (1) interventions push the exchange rate in the desired direction or (2) interventions reduce exchange rate volatility. The desired direction in the case of Japan is a depreciation of the yen most of the time.¹ In terms of equation (2), an appreciation of the yen as a consequence of interventions amounts to a positive coefficient α_1 .

¹On 10-Apr-1998, the Japanese monetary authorities sold \$20.4bn. As can be seen in Figure 1 this is a rare and by far the largest instance where an intervention to appreciate the yen against the dollar was undertaken. Deleting this “outlier” does not substantially change our results.

Judging by these standards, interventions have not done well on the first segment between 1991 and 1995: The coefficient α_1 of interventions in the returns equation (2) is significantly negative. It has the wrong sign in the sense that dollar purchases show a significant correlation with appreciations of the yen. The coefficient β_4 of interventions in the log realized volatility equation (3) is significantly positive. That is, dollar purchases exhibit a significant correlation with increases in realized volatility of the yen-to-dollar exchange rate. Hansen's (1982) J -statistic for the overidentifying restrictions is 0.005, which is less than the critical value for any common confidence level and three degrees of freedom.² Therefore, we cannot reject the null hypothesis that the overidentifying restrictions hold.

On the second segment between 1995 and 1998, interventions have a significantly positive coefficient $\alpha_1 = 0.0056$ in the return equation (2). During this period, interventions seem to have had the desired effect on the yen-to-dollar exchange rate. The standard deviation of daily returns on the exchange rate in this period is 0.0069, so the "marginal" influence of one billion dollars volume of intervention purchases amounts to a movement of roughly 4/5 of a standard deviation. The coefficient β_4 of the intervention variable in equation (3) is not significantly different from zero. In this period, interventions can be considered successful in the sense of influencing the returns on the exchange rate. The J -statistic again indicates validity of the overidentifying restrictions.

The coefficient of interventions α_1 in equation (2) on the third segment from 1998 to 2004 is not significantly different from zero. Therefore, interventions did not have the desired effect on the returns. The coefficient β_4 of interventions in equation (3) is highly significantly negative, however. Interventions show significant negative correlation with realized volatility. They can be considered

²There are eight slope coefficients in model (2) to (4). We count the weekly and monthly averages over the lags of log realized volatilities as one variable, respectively. Equation (2) has one right-hand side variable (besides the constant) and two instruments. Equation (3) has four right-hand side variables and five instruments. Equation (4) has three right-hand side variables and four instruments. Thus there are three overidentifying conditions.

successful in the sense of reducing volatility of the exchange rate. We do not reject the null hypothesis of validity of the overidentifying restrictions.

Estimated on the entire sample, we obtain a return coefficient α_1 of interventions of essentially zero and a highly significantly negative effect on realized volatility of $\beta_4 = -0.0595$. The J -statistic on the entire sample is 0.016, failing to reject the null hypothesis. We do not include a table for the total sample for brevity.

In summary, in our model we observe a change in the success of interventions over time. During the period 1991 to 1995 when Japanese authorities pursued a strategy of frequent relatively small interventions, interventions coincided with movements of the yen-to-dollar rate in the “wrong” direction and with increases in realized volatility. During the period from 1995 to the deregulation of the exchange rate market in 1998, interventions were successful in influencing the exchange rate in the right direction and had no influence on realized volatility. After the deregulation in 1998 until after the last episode of interventions in 2004, interventions had no significant influence on the returns of the exchange rate but have decreased realized volatility.

4 Robustness Analysis

In this section, we relax the structural assumptions that we have made on the model for system y so far. We consider other financial variables as determinants in equations (2) through (4) such as stock index prices and interest rates; we consider unrestricted vector autoregressions, and alternative change-points. Most of the results reported in Section 3.4 remain visible in these alternative specifications. We keep the reporting to a qualitative minimum in order not to clutter the paper with tables.³

³The corresponding author’s web site www.bus.lsu.edu/hillebrand provides an EViews work file for download that contains all data and results.

4.1 Vector-Autoregressive Specifications

It is possible that the results reported in Section 3.4 depend on the specific structural form of equations (2) through (4), the GMM estimation method, or the choice of instruments. In order to assess this possibility, we estimate different vector autoregressive specifications for the system y . If we drop the modulus operator $|\cdot|$ from the intervention variable I_t in equation (3), the system can be written as⁴

$$Ay_t = \omega + \Phi_1 y_{t-1} + \sum_{j=2}^5 \Phi^{(w)} y_{t-j} + \sum_{j=6}^{21} \Phi^{(m)} y_{t-j} + \varepsilon_t, \quad (5)$$

where $y_t = (r_t, \log \sigma_t, I_t)$, ε_t is a three-dimensional zero-mean process and

$$A = \begin{bmatrix} 1 & 0 & -\alpha_1 \\ 0 & 1 & -\beta_4 \\ -\gamma_2 & 0 & 1 \end{bmatrix}, \quad \Phi_1 = \begin{bmatrix} 0 & 0 & 0 \\ 0 & \beta_1 + \frac{\beta_2}{5} + \frac{\beta_3}{21} & 0 \\ \gamma_3 & 0 & \gamma_1 \end{bmatrix},$$

$$\Phi^{(w)} = \begin{bmatrix} 0 & 0 & 0 \\ 0 & \frac{\beta_2}{5} + \frac{\beta_3}{21} & 0 \\ 0 & 0 & 0 \end{bmatrix}, \quad \Phi^{(m)} = \begin{bmatrix} 0 & 0 & 0 \\ 0 & \frac{\beta_3}{21} & 0 \\ 0 & 0 & 0 \end{bmatrix}.$$

Dropping all structural constraints, this specification suggests an unrestricted VAR of order 21. Figure 4 graphs the impulse responses of the returns r_t to interventions I_t in the top row of panels and the impulse responses of log realized volatility $\log \sigma_t$ to interventions I_t in the bottom row of panels. There are three panels in each row for the three sub-periods of data.

The first panel (a) shows the significantly negative response of returns to interventions, confirming the “wrong” sign of the intervention effect. We cannot confirm the significantly positive influence on volatility in panel (b). For the period 1995 to 1998, both impulse responses are not significantly different from zero, so we cannot conclude any influence of interventions from the VAR(21) for this period. For the last period, the response of returns in panel (e) is again not significantly different from zero, which corresponds to the findings

⁴Figure 1 shows that by far the largest part of interventions has positive sign.

in the structural equation. The response of volatility in panel (f), however, confirms the negative impact of interventions on volatility up to lag five. This confirms the finding that in the last regime 1998 to 2004, interventions seem to have reduced volatility. In contrast to the structural model, which registers the influence only contemporaneously, the VAR(21) estimation indicates a dynamic negative influence up to five days after the intervention.

FIGURES 4, 5 ABOUT HERE

Even though there are $3 \times 3530 = 10590$ observations to support the estimation, the VAR(21) specification may suffer from overfitting, in particular on the segment 1995 to 1998, where there are only $3 \times 736 = 2208$ observations (to support the 189 autoregressive coefficients to be estimated). We choose a VAR(3) specification as a more parsimonious alternative. (Lag selection criteria such as likelihood ratio tests are not valid due to the non-standard stochastic structure of the intervention time series. These standard tests indicate an optimal lag at about 20.) Figure 5 shows the impulse responses for the VAR(3) specification. The organization is the same as in the previous graph, each column shows the responses to interventions of the returns in the top panel and of the volatility in the bottom panel. There are three columns for the three sub-periods.

Panel (a) of Figure 5 confirms the negative influence of interventions on returns in the first segment 1991 to 1995. In contrast to the VAR(21) specification, the VAR(3) confirms in panel (b) that interventions increased volatility in that period. Panel (c) shows a largely positive response of returns to interventions in the period 1995 to 1998, but it is not strong enough to be significant. As in the case of the structural model, the volatility response is insignificant in the middle period as shown in panel (d). The response of returns to interventions during 1998 to 2004 in panel (e) is basically zero, as found in the other specifications. Volatility in this period, however, shows a very strong response in panel (f). As found in both, the structural model and the VAR(21), there is a strong decrease in volatility in reaction to the intervention. In this VAR(3) specification the

response is significantly negative beyond lag ten.

In summary, the unrestricted vector autoregressive specifications largely confirm the findings reported in Section 3.4. The only exception is the successful influence of interventions on returns in the 1995 to 1998 period, which is only found in the structural model. The VAR specifications strongly confirm the smoothing influence of interventions during the 1998 to 2004 period and indicate a decrease in volatility after interventions for several days.

4.2 More General Structural Specifications

In order to capture the influence of other asset markets on exchange rate and interventions, we study the influence of returns on the daily Nikkei 300 index in equations (2) to (4). We add daily returns on the Nikkei to equation (2) and its instruments, we add the log of volatility obtained from daily returns on the Nikkei to equation (3) and its instruments, and we add daily returns on the Nikkei to equation (4) and its instruments. Across all three segments, the sign and significance of the estimates of the parameters of interest, α_1 and β_4 , remain the same as reported for the simple specification.⁵

We also estimated a version of the structural model that keeps equations (2) and (3), but adds contemporaneous log realized volatility and its first lag to equation (4). On all segments the signs of the intervention parameters remain the same as in the simple specification but the significance is only at the 10% level.

Interest rates are another potentially important determinant in the system y . To explore this influence, we include the change in the Japanese uncollateralized overnight interbank rate (Mutan) $i_{t,\text{mutan}} - i_{t-1,\text{mutan}}$ as well as the interest rate differential with the US Federal Funds Rate $i_{t,\text{FFR}} - i_{t,\text{mutan}}$ at concurrent values and at lag one in all three equations of system (2) through (4). Since the Japanese Mutan rate is available only after 11-Apr-1996, we estimate the

⁵We also included the Dow Jones Industrial Average in addition to the Nikkei and replacing the Nikkei. The results were very similar.

segment on two sub-periods: (1) 11-Apr-1996 to 1-Apr-1998 and (2) 2-Apr-1998 to 27-Oct-2004. Note that the Mutan rate became almost zero in early 1999 and did not change substantially until 2004.

On the first segment 1996–1998, we obtain no influence of interventions on returns (α_1 is not significantly different from zero). The volatility coefficient β_4 of interventions, however, is significantly negative at the 1% level. The concurrent change in the Mutan rate and the concurrent interest differential have significantly negative influence on returns, and the lagged interest differential has significantly positive influence on volatility. All other interest rate coefficients are not significantly different from zero.

On the second segment 1998–2004, the result is qualitatively the same as in the simple specification: Interventions have no influence on returns and decrease volatility at the 1% significance level. The interest rate coefficients are at best significant at the 10% level in this segment.

In summary, adding the Nikkei 300 index returns and volatility to the model does not change the results with respect to interventions qualitatively. Including volatility in the reaction function does not change the sign of the coefficients of interest but weakens their significance. Adding changes in the Japanese overnight interbank interest rate and the interest rate differential with the Federal Funds Rate, we obtain an insignificant influence of interventions on returns and a significant reduction of volatility by interventions on both sub-segments 1996–1998 and 1998–2004. In the case of interest rates, the period 1991–1996 cannot be studied due to data unavailability.

4.3 Alternative Change-Points

We apply the change-point detector statistic proposed in Bai (1994, 1997) to the series of logarithmic realized volatilities displayed in Figure 3. The asymptotic results about the statistic hold as long as realized volatility can be described by a linear time series model such as equation (3), thus our model specification

is in line with the assumptions made in Bai (1994). The statistic converges in distribution under the null hypothesis to a standard Brownian Bridge. We choose the VARHAC estimator of Den Haan and Levin (1997) as estimator of the variance of the log realized volatility series, which is necessary to compute the statistic. At the 99% confidence level, the detector returns two significant change-points: 1-Mar-1995 and 9-Jun-2000. The corresponding test statistics are 2.23 and 2.47, respectively, indicating p-values of essentially zero (non-zero entries in fifth decimal place).

Adopting these alternative change-points, we repeat the estimation of model (2) through (4). We only report the qualitative findings for the main parameters of interest, α_1 and β_4 . On the first segment 1991 to 1995, we obtain a negative α_1 significant at the 5% level and an estimate of β_4 of essentially zero. On the second segment 1995 to 2000, the estimate of α_1 is not significantly different from zero and the estimate of β_4 is significantly negative at the 5% level. On the last segment 2000 to 2004, the estimate of α_1 is not significantly different from zero and the estimate of β_4 is significantly negative at the 1% level.

In summary, using the change-points indicated by a statistical change-point detector, the results of the structural model are confirmed for the early period 1991–1995 (except for the increase in volatility) and for the late period 2000–2004. The intermediate period 1995–2000 already picks up a significantly negative impact of interventions on volatility but does not show any success of interventions in the return equation.

5 Conclusion

This study examines the interrelations of returns and realized volatility of the yen-to-dollar exchange rate with interventions of the Japanese monetary authorities. The concept of realized volatility allows us to treat volatility as a variable observed with measurement error and enables us to employ a simultaneous equations model for returns, realized volatility, and interventions. This

resolves the endogeneity problem that plagued earlier approaches to measure the success of interventions.

The sample period considered is April 1991 through October 2004. The literature has identified two structural breaks in the conditions of interventions during that period: In June of 1995, the Japanese authorities switched from a policy of frequent and relatively small interventions to infrequent and relatively large interventions (Itô and Yabu 2004). In April of 1998, the foreign exchange market in Japan was largely deregulated (Itô and Melvin 1999). Estimating our system of equations on the resulting three sub-segments of our sample, we find that during the period 1991 to 1995, interventions were unsuccessful. The returns to the yen-to-dollar rate responded significantly negative and volatility was increased. During the period 1995 to 1998, exchange rate volatility was unaffected by interventions but we find some evidence of successful influence on returns. During the later period 1998 to 2004, returns are unaffected by interventions but we find strong evidence of a reduction in volatility.

The robustness of the results is studied in unrestricted vector autoregressions. We also consider more general versions of the structural model that include other financial variables such as stock index prices and interest rates. Alternative change-point choices according to a statistical change-point detector are examined. These checks largely confirm the findings from the structural model. The most robust result is the significant reduction in volatility that interventions achieved in the more recent episodes.

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Table 1: SUMMARY STATISTICS FOR INTERVENTIONS.

	Bank of Japan	Federal Reserve
Intervention Days	335	22
Days of Dollar Purchases	304	18
Coinciding		18
Volume Purchases (\$ Bn)	579.2	7.5
Days of Dollar Sales	31	4
Coinciding		3
Volume Sales (\$ Bn)	-37.8	-1.1

“Coinciding” means that interventions of the Japanese authorities and of the Federal Reserve occur on the same day and have the same sign.

Table 2: ESTIMATION OF MODEL (2) THROUGH (4) FOR SAMPLE 2-APR-1991 THROUGH 1-JUN-1995 (OBSERVATIONS 1 THROUGH 1084).

LHS	coeff.	RHS	estimate	std. err.	t-stat.	prob.
r_t	α_0	const	-0.0003	0.0001	-1.94	0.053
	α_1	I_t	-0.0036	0.0013	-2.77	0.006
$\log \sigma_t^2$	β_0	const	-0.3181	0.0760	-4.12	0.000
	β_1	$\log \sigma_{t-1}^2$	0.3286	0.0422	7.78	0.000
	β_2	$\log \sigma_{t-1,w}^2$	0.4006	0.0521	7.69	0.000
	β_3	$\log \sigma_{t-1,m}^2$	0.1299	0.0498	2.61	0.009
I_t	β_4	I_t	0.1172	0.0533	2.20	0.028
	γ_1	I_{t-1}	0.4176	0.1086	3.85	0.000
	γ_2	r_t	-23.36	22.22	-1.05	0.293
	γ_3	r_{t-1}	-48.41	30.86	-1.57	0.117

The symbols $\log \sigma_{t-1,w}^2$ and $\log \sigma_{t-1,m}^2$ stand for the averages of log realized volatility over 5 lags and over 21 lags as specified in equation (3). The J -statistic for validity of the overidentifying restrictions is 0.0053.

Table 3: ESTIMATION OF MODEL (2) THROUGH (4) FOR SAMPLE 2-JUN-1995 THROUGH 1-APR-1998 (OBSERVATIONS 1085 THROUGH 1820).

LHS	coeff.	RHS	estimate	std. err.	t-stat.	prob.
r_t	α_0	const	0.0004	0.0002	1.55	0.122
	α_1	I_t	0.0056	0.0015	3.68	0.000
$\log \sigma_t^2$	β_0	const	-0.3550	0.0915	-3.88	0.000
	β_1	$\log \sigma_{t-1}^2$	0.3780	0.0477	7.92	0.000
	β_2	$\log \sigma_{t-1,w}^2$	0.3011	0.0718	4.19	0.000
	β_3	$\log \sigma_{t-1,m}^2$	0.1493	0.0694	2.15	0.032
	β_4	I_t	0.0109	0.0413	0.26	0.791
I_t	γ_1	I_{t-1}	0.0543	0.0410	1.33	0.185
	γ_2	r_t	6.823	15.99	0.43	0.670
	γ_3	r_{t-1}	-13.28	18.82	-0.71	0.480

The symbols $\log \sigma_{t-1,w}^2$ and $\log \sigma_{t-1,m}^2$ stand for the averages of log realized volatility over 5 lags and over 21 lags as specified in equation (3). The J -statistic for validity of the overidentifying restrictions is 0.027.

Table 4: ESTIMATION OF MODEL (2) THROUGH (4) FOR SAMPLE 2-APR-1998 THROUGH 27-OCT-2004 (OBSERVATIONS 1821 THROUGH 3530).

LHS	coeff.	RHS	estimate	std. err.	t-stat.	prob.
r_t	α_0	const	-0.0004	0.0001	-3.50	0.001
	α_1	I_t	-0.0001	0.0002	-0.30	0.767
$\log \sigma_t^2$	β_0	const	-0.1643	0.0568	-2.90	0.004
	β_1	$\log \sigma_{t-1}^2$	0.3641	0.0500	7.28	0.000
	β_2	$\log \sigma_{t-1,w}^2$	0.3386	0.0665	5.09	0.000
	β_3	$\log \sigma_{t-1,m}^2$	0.2184	0.0482	4.53	0.000
I_t	β_4	I_t	-0.0655	0.0204	-3.21	0.001
	γ_1	I_{t-1}	0.2969	0.0647	4.59	0.000
	γ_2	r_t	-11.95	58.75	-0.20	0.839
	γ_3	r_{t-1}	-30.47	54.00	-0.56	0.573

The symbols $\log \sigma_{t-1,w}^2$ and $\log \sigma_{t-1,m}^2$ stand for the averages of log realized volatility over 5 lags and over 21 lags as specified in equation (3). The J -statistic for validity of the overidentifying restrictions is 0.025.

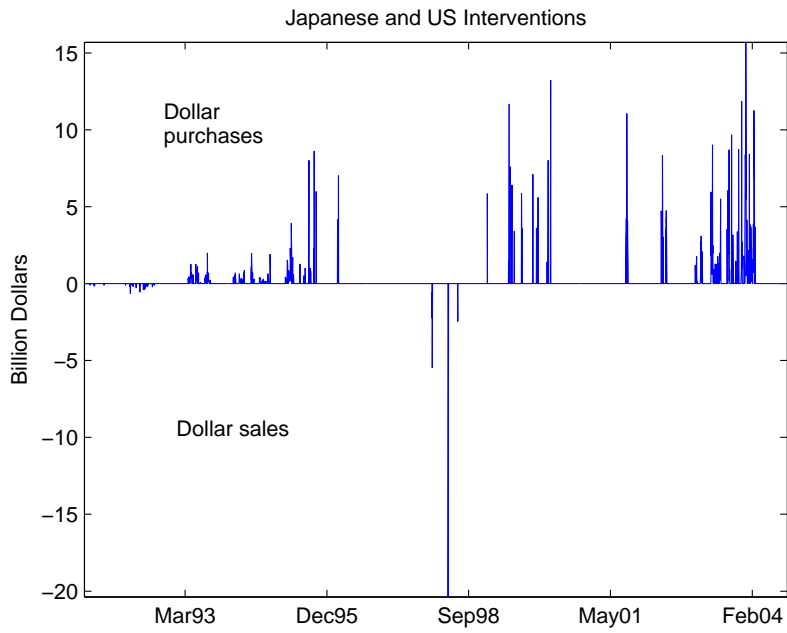


Figure 1: Daily foreign exchange interventions by Japanese and U.S. authorities in the yen-to-dollar market during 1991 to 2004 measured in Billion Dollars. Dollar purchases (yen sales) have positive sign; dollar sales (yen purchases) have negative sign. Source: Japanese Ministry of Finance, <http://www.mof.go.jp/english/e1c021.htm>.

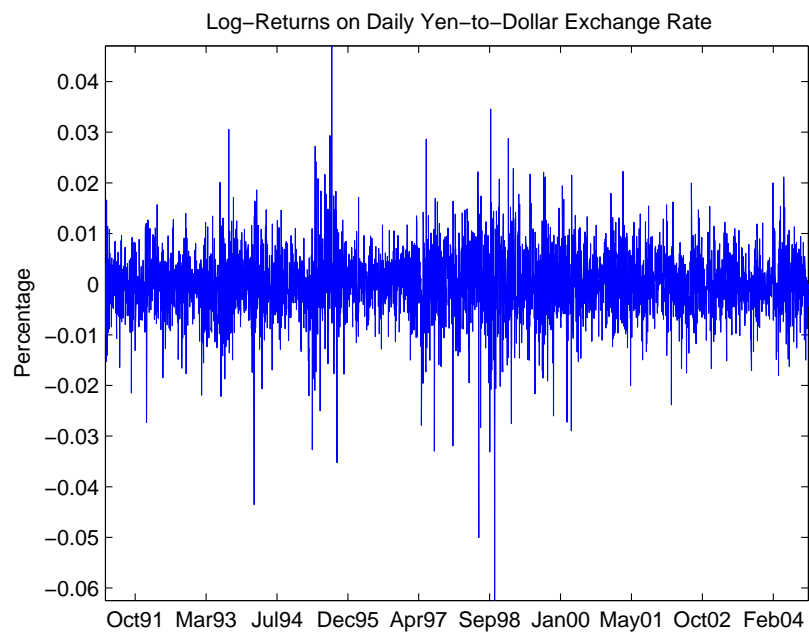


Figure 2: Daily log returns on the yen-to-dollar exchange rate from 1991 to 2004. Source: Olsen Financial Technology.

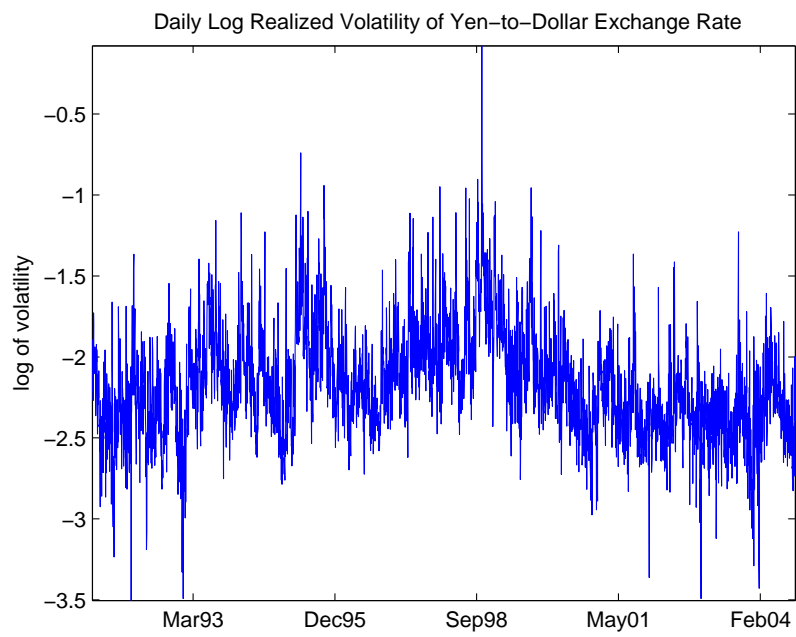


Figure 3: Daily log realized volatility of the yen-to-dollar exchange rate from 1991 to 2004. Source: Olsen Financial Technology.

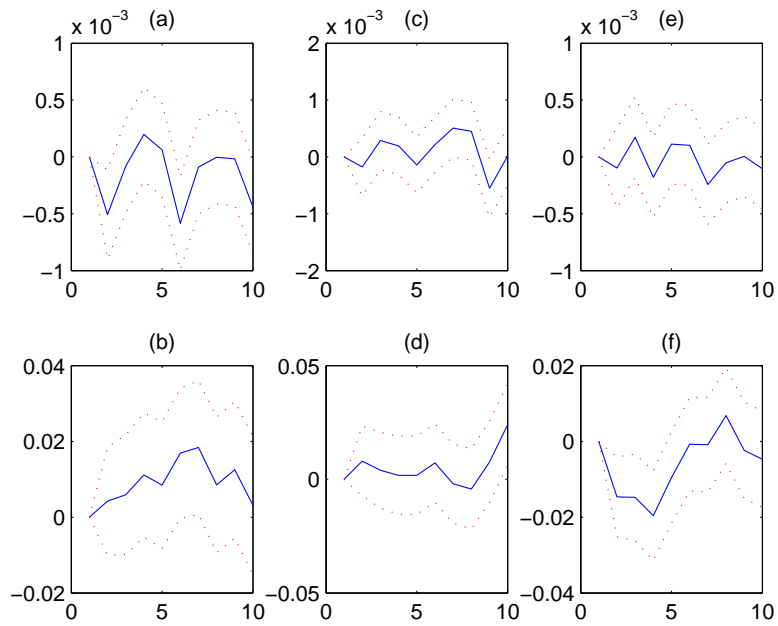


Figure 4: VAR(21) estimation. Orthogonalized impulse responses of returns on the yen-to-dollar rate to interventions (top row: panels (a) 1991–1995, (c) 1995–1998, (e) 1998–2004); orthogonalized impulse responses of log realized volatility to interventions (bottom row: panels (b) 1991–1995, (d) 1995–1998, (f) 1998–2004)

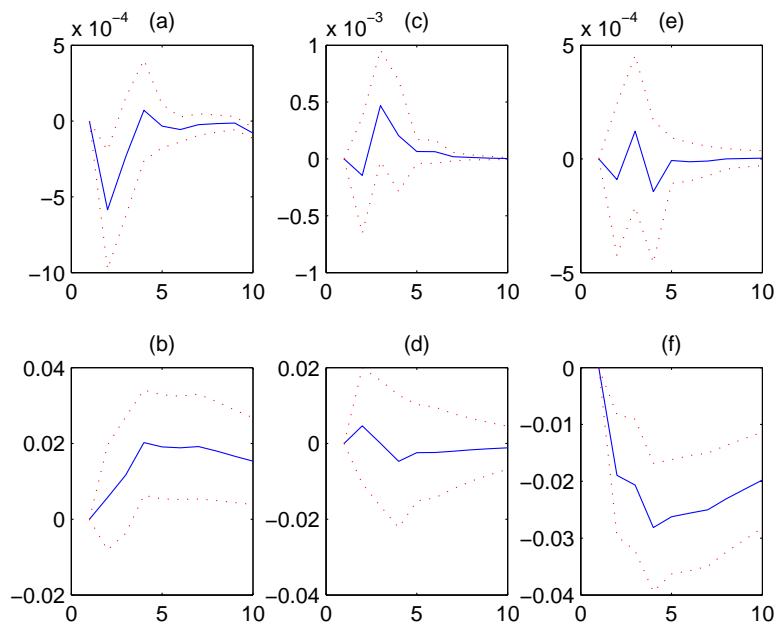


Figure 5: VAR(3) estimation. Orthogonalized impulse responses of returns on the yen-to-dollar rate to interventions (top row: panels (a) 1991–1995, (c) 1995–1998, (e) 1998–2004); orthogonalized impulse responses of log realized volatility to interventions (bottom row: panels (b) 1991–1995, (d) 1995–1998, (f) 1998–2004)