

Forti Grazzini, Caterina; Rieth, Malte

Conference Paper

Interest Rates and Exchange Rates in Normal and Crisis Times

Beiträge zur Jahrestagung des Vereins für Socialpolitik 2017: Alternative Geld- und Finanzarchitekturen - Session: Exchange Rates, No. C14-V2

Provided in Cooperation with:

Verein für Socialpolitik / German Economic Association

Suggested Citation: Forti Grazzini, Caterina; Rieth, Malte (2017) : Interest Rates and Exchange Rates in Normal and Crisis Times, Beiträge zur Jahrestagung des Vereins für Socialpolitik 2017: Alternative Geld- und Finanzarchitekturen - Session: Exchange Rates, No. C14-V2, ZBW - Deutsche Zentralbibliothek für Wirtschaftswissenschaften, Leibniz-Informationszentrum Wirtschaft, Kiel, Hamburg

This Version is available at:

<https://hdl.handle.net/10419/168281>

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.

Interest Rates and Exchange Rates in Normal and Crisis Times

Caterina Forti Grazzini¹ and Malte Rieth²

February 17, 2017

JEL classification number: E44, F3, G1

¹Humboldt-Universität zu Berlin and DIW Berlin, Macroeconomic Department.
E-mail: cfortigrazzini@diw.de.

²DIW Berlin, Macroeconomic Department. E-mail: mrieth@diw.de.

Interest Rates and Exchange Rates in Normal and Crisis Times

Abstract

The paper studies the empirical relation between the US-Dollar/Euro exchange rate and US and euro area interest rates during normal and crisis times. We describe each asset price within a multifactor model and identify the causal contemporaneous relations through heteroskedasticity. We find that US interest rates and macroeconomic conditions dominate exchange rate and interest rate movements before and during the global financial crisis, while this pattern sharply reverses during the European debt crisis with euro area developments playing the leading role. The announcement of Outright Monetary Transactions by the European Central Bank seems to have restored pre-crisis normality, albeit with now both regions being similarly important for international asset price formation.

Keywords: International financial markets; global financial crisis; European debt crisis; identification through heteroskedasticity; United States; Euro area.

JEL classification numbers: E44; F31; G1.

1 Introduction

The integration of global financial markets has increased substantially during the last two decades with US markets, investors, and monetary policy playing a main role in the determination of domestic and international asset prices (Rey, 2015). This integration process, however, also saw periods of sharp reversals. During the global financial crisis investors repatriated funds to the US (Fratzscher, 2009), for example, and in the European debt crisis local non-US developments seem to have mainly influenced the pricing of domestic asset prices and bilateral exchange rates (Ehrmann et al., 2014). In addition, the unconventional measures adopted by the US Federal Reserve and subsequently by other major central banks in response to these crises have had large impacts on both domestic and international financial markets (Rogers et al., 2014), although the debate regarding the magnitude of these effects is still ongoing. A case in point reflecting these developments is the USD-EUR exchange rate (see Figure 1). It depreciated sharply after the fall of Lehman Brothers, with hitherto unseen increases in volatility during the following crises episodes.

While several recent papers investigate whether the behavior of exchange rates and interest rates is affected by the zero lower bound in the US (Swanson and Williams, 2014a; Stavrakeva and Tang, 2016) or by specific crisis episodes (Fratzscher, 2009; Ehrmann et al., 2014), a structural analysis that systematically compares exchange rate formation during crisis and non-crisis periods is large missing in the literature. In this paper, we aim to fill this gap by assessing whether and how the recent two crisis episodes have changed the international transmission of interest rate and news shocks and in particular their effects on the exchange rate. Understanding the dissipation of interest rate conditions across countries and over time is important for policymakers and practitioners alike as domestic interest rates play a key role for intertemporal decisions of households, firms, and governments. Just as essential is a better understanding of the behavior of the nominal exchange rate because this price determines foreign demand for home goods if prices are sticky.

We focus on the USD-EUR exchange rate and US and euro area interest rates and use daily data since the introduction of the euro. We split the sample into four subperiods following what the literature has identified as the main game-changers for the US and the euro area. We concentrate our attention on these areas as they are the two largest countries in the world with flexible exchange rates and are thus expected to be less affected by external developments than small open economies. A main difficulty in identifying

the causal effects between interest rate changes and exchange rate movements is the endogeneity of these asset prices even at the daily frequency. We address this challenge by modeling each variable with a multifactor model and use the heteroskedasticity in the data to identify the contemporaneous impacts. In a nutshell, this identification strategy exploits the fact that periods of, say, higher interest rate volatility contain additional information on the response of the exchange rate to interest rate changes as the latter are more likely to occur. These changes in the volatility of interest rate shocks can then be used as a “probabilistic instrument” (Rigobon, 2003) to trace out the impact of interest rate shocks on the exchange rate. Similarly, the framework allows quantifying the strength of contemporaneous interest rate spillovers across the Atlantic.

We find significant changes in the structural relations among the endogenous variables across periods. In general, US economic conditions dominate exchange rate and interest rate formation in tranquil times, while during turbulent times asset prices become more sensitive to developments in the economy where the crisis originates. In particular, in the pre-crisis period, only US interest rate shocks have a significant effect on the exchange rate, but not those from the euro area. This asymmetry increases during the global financial crisis, but sharply reverses during the European crisis, with euro area interest rate shocks having the largest impact on the exchange rate. Interestingly, the announcement of the Outright Monetary Transactions (OMTs) programme by the European Central Bank (ECB) seems to have re-established pre-crisis normality in the foreign exchange market in the sense that US interest rates shocks re-gained their importance, although the exchange rate remains more sensitive to euro area interest rate shocks than before the crises. We find similar patterns for the bidirectional spillovers between US and euro area interest rates, with US rates being more important in all periods except for the euro area crisis sample.

These findings are supported by forecast error variance decompositions which show that US interest rate shocks typically explain a larger fraction of exchange rate and interest rates variability than euro area interest rate shocks, but that the latter caught up with their US counterparts. Finally, news regressions confirm these patterns. They show that the bilateral exchange rate became increasingly more sensitive to euro area macroeconomic data surprises, which now play a similarly important role as news shocks from the US.

Our results highlight the changing nature of international financial linkages both across crisis and non-crisis periods and with respect to the underlying financial integration process between the US and the euro area. First,

they indicate that the relation between interest rates and exchange rates may change during crisis episodes, providing a potential explanation for why uncovered interest parity may not hold on average (Engel, 2014). Specifically, they suggest that the sensitivity of the exchange rate to the fundamentals of the origin country of the crisis increases. Second, they point to progressing financial integration and an increased importance of euro area economic conditions for international asset pricing as asset prices now respond more to shocks coming from this region than in the pre-crisis period. This finding complements earlier work by Ehrmann and Fratzscher (2005a) who document an increasing interdependence between the two regions starting with the introduction of the euro. Finally, in terms of policy implications, our findings support the notion that the ECB restored a normal functioning of exchange rate pricing through the announcement of OMTs, while retaining some influence of euro area interest rate conditions on the exchange rate.

Our paper relates to several strands of literature. Generally, it connects to empirical studies of the relationship between exchange rates and interest rates. Ehrmann et al. (2011) use a vector autoregressive (VAR) model estimated on pre-crisis data and find that both US and euro area interest rate shocks are important for USD/EUR exchange rate developments. A long standing history of papers analyzes the relationship within the framework of the uncovered interest rate parity (UIP), which states that (lagged) short and long interest rates are fundamental in determining the exchange rate. UIP is a cornerstone of international finance, constituting an important building block of many exchange rate determination theories. The empirical evidence supporting UIP is mixed, however. While several papers find evidence in favor of this relation (Baillie and Bollerslev, 2000; Chinn and Meredith, 2004), others document its failure (Bekaert et al., 1997; Engel, 1996; Bekaert et al., 2007).

Another, more specifically related literature analyzes whether the behavior of exchange rates changes during particular time periods. Focusing on the global financial crisis, Fratzscher (2009) finds that this episode has triggered sharp and unexpected currency movements, with domestic macroeconomic fundamentals and financial exposure of individual countries playing a key role in the transmission of US shocks. Concentrating on the European crisis, Ehrmann et al. (2014) present evidence that the euro mainly danced to its own tune and that fundamentals and policy decisions possess only little explanatory power for exchange rate fluctuations. Swanson and Williams (2014b) investigate the effect of the zero lower bound on asset price formation using an event study design. They show that interest rates are partially constrained during this period, while exchange rates are not. Sim-

ilarly, Stavrakeva and Tang (2016) show that the contemporaneous relation between yields at different maturities and exchange rates has changed after the zero lower bound was hit in the US, based on a model with VAR-based expectations of short-run yields and inflation. Several other papers analyze whether there are differences between the effects of conventional and unconventional monetary policy shocks on exchange rates (Glick et al., 2013; Kiley, 2013; Glick et al., 2015; Ferrari et al., 2016). The main difference between these papers and our work is that we do not study the relationship between exchange rate and interest rates conditional on a particular shock (in this case, a monetary policy shock), but more generally study the reaction of the exchange rate to an interest rate shock, regardless of what has caused it.

Finally, another strand of literature linked to our work assesses the effects of macroeconomic news and exchange rates and interest rates. Andersen et al. (2003) show that announcement surprises produce conditional mean jumps in a broad set of foreign exchange rates. Ehrmann and Fratzscher (2005b) address this issue, considering daily the USD-EUR exchange rate from 1999 to 2003. They find that news about fundamentals can explain both the direction and the magnitude of daily and monthly exchange rate developments and that US news have a larger impact than the euro area. More recent papers analyse the impact of macroeconomic news on German interest rates and USD-EUR exchange rate (Swanson and Williams, 2014b) as well as on the Treasury yields (Swanson and Williams, 2014a), and measure whether and to what extent the latter have been affected by the zero lower bound on nominal interest rates. They find that the exchange rate has been essentially unaffected by this bound, while the German and the Treasury yields start being constraint only after 2011.

The paper is organized as follows. In Section 2 we outline the empirical model and describe the data. Section 3 contains the main results, before the last section concludes.

2 Methodology and data

In this section we first describe the model set-up and the data. Then, we discuss the problem of identification in the context of structural vector autoregressive (VAR) models and how identification through heteroskedasticity gives a suitable solution.

2.1 Model specification and data

The structural VAR model is

$$Ay_t = \tilde{c} + \tilde{A}_1 y_{t-1} + \dots + \tilde{A}_p y_{t-p} + \tilde{\Gamma} x_t + \epsilon_t, \quad (1)$$

where y_t and x_t are the vectors of endogenous and exogenous variables, \tilde{c} is a vector of constants, \tilde{A}_i and $\tilde{\Gamma}$ with $i = 0, \dots, p$ are coefficient matrices and ϵ_t is a vector of structural shocks with the following regime-dependent diagonal variance matrix in regime k :

$$\Sigma_{\epsilon,k} = E(\epsilon_t \epsilon_t') = \begin{Bmatrix} \sigma_k^1 & 0 & 0 \\ 0 & \sigma_k^2 & 0 \\ 0 & 0 & \sigma_k^3 \end{Bmatrix}.$$

The focus of the paper is on the impact matrix A , which contains the contemporaneous effects of structural shocks on the endogenous variables. The vector y_t includes the nominal exchange rate, measured as the amount of US-Dollars per one euro such that an increase in the variable reflects a depreciation of the US-Dollar, the risk-free rate in the euro area, approximated through the two-year rate on German government bonds, and the two-year rate on US Treasury bonds:

$$y_t = \begin{Bmatrix} e_t \\ r_t^{EU} \\ r_t^{US} \end{Bmatrix}.$$

We collect data from Datastream and Bloomberg for the period January 1st 2000 to November 3rd 2016. We split the data into four subsamples: the pre-crisis period, the global financial crisis period, the European sovereign crisis period and the post-OMT period. The first subsample, the pre-crisis period, runs from the beginning of the sample until August 9, 2007, when the large French bank BNP Paribas temporarily halted redemptions from three of its funds that held assets backed by US subprime mortgage debt. This event has been seen by many commentators as the trigger of the global financial crisis (among the others, the former ECB President Trichet, 2010 and Cecchetti, 2008), and we use it as starting point of our second subsample. The second subsample covers from 10 August 2007 until 1 September 2009, when the newly elected Greek government announced for the first time that there could be problems with the Greek government debt data. Then the European sovereign crisis starts and it goes on until 30 September 2012, that is after President Draghis “Whatever it takes” speech on July 26th

2012, after the ECB’s official announcement of the OMT at the beginning of August 2012 and after the announcement of the implementation details for the OMT program in September of the same year. Finally, the last subsample, the post-OMT, runs from 1 October 2012 until the end of the sample, on 3 November 2016. Moreover, as we will formally show in the main analysis, the data support this sample split, as structural relations among the endogenous variables change significantly across subsamples.

To account for macroeconomic news shocks, we include the unexpected or “surprise” component of releases of economic indicators for both the US and the euro area as exogenous control variables. In fact, several papers have shown that these news are important for the development of exchange rates (among the others, Andersen et al. 2003 and Swanson and Williams 2014b). The complete list of the included indicators can be found in table 6. The surprise component of each announcement is computed as the actual realization of the economic indicators minus the financial market’s expectations from few days before. The data are obtained from Bloomberg, which conduct a survey about financial markets institutions and professional forecasters about their expectations of future data releases. The data provider collects informations from the survey participants up to the night before the release. Thus, the forecast should reflect all information available up to few hours before the release, which should allow us to capture only the unexpected component, i.e. the only part that should have an effect on financial markets.

The data are at a daily frequency. Following Ehrmann et al. (2011), we construct two-day windows because US shocks may occur after the closing of European markets, thus affecting the latter only on the next business day. To avoid losing information on macroeconomic releases, all news that would happen on an excluded day are moved forward and reported as if they occurred on the next business day. Finally, to account for the non-stationarity of data we estimate the model in first differences of interest rates and log differences of the exchange rate.¹ All variables are standardized prior to estimation, as this enable us to compare the relative effects of variables having different units of measurement. Finally, in the analysis of each subsample we include one lag of the endogenous variables, based mainly on the BIC criteria.

¹We perform augmented Dickey-Fuller and Phillips-Perron tests on the levels of the interest rates and on the first difference of the exchange rate. The tests do not reject the null hypothesis of a unit root.

2.2 Identification through heteroskedasticity

For estimation we pre-multiply the structural VAR model (1) by A^{-1}

$$y_t = A^{-1}\tilde{c} + A^{-1}\tilde{A}_1y_{t-1} + A^{-1}\tilde{A}_2y_{t-2} + A^{-1}\tilde{\Gamma}x_t + A^{-1}\epsilon_t \quad (2)$$

and re-write it as

$$y_t = c + A_1y_{t-1} + A_2y_{t-2} + \Gamma x_t + u_t \quad (3)$$

where the vector of reduced form shocks $u_t = A^{-1}\epsilon_t$ is related to the structural shocks through matrix A . The parameters of this reduced form can be consistently estimated by ordinary least squares. The coefficients of the structural model can be recovered only if the contemporaneous impact matrix is identified, however. Hence, we need to make some assumptions on how to recover the elements of this matrix from the reduced form estimates. Using (2) and (3) it follows that $\Sigma_u = A^{-1}\Sigma_\epsilon(A^{-1})'$. Thus, Σ_u can in principle be estimated through this relation, if the number of unknown parameters equals the number of linearly independent equations. This is however not the case and we thus need additional information to identify the model.

A common practice in the structural VAR literature is to impose additional restrictions in the form of zero restrictions or sign restrictions. However, as exchange rates and government interest rates react simultaneously to each other, it is impossible to impose short run-restrictions in the form of zero restrictions. One alternative option would be employing sign restrictions, which do allow for contemporaneous effects among the variables. The side effect of this methodology in the context of our analysis is that it constrains the contemporaneous response of being of a (ex-ante determined) particular sign. Unfortunately, there is no consensus in the literature yet on the bidirectional causality between asset price relationships, and it is thus hard to come up with a robust sign scheme that would allow the identification of shocks. For example, Fratzscher (2009) finds that while before the global financial crisis, a negative US shock would lead to a depreciation of the US dollar against foreign currencies, in period of crisis this is no longer true and that negative news about the US economy would make the dollar appreciate. Also Rogers et al. (2016) document that during the period of crisis, monetary policy easing done by the ECB leads to a euro appreciate, in contrast to what standard theory would predict.

To identify the shocks we thus use an alternative methodology, known as identification through heteroscedasticity, developed by Sentana and Fiorentini (2001) and Rigobon (2003). This methodology exploits the fact that financial variables are generally found to be heteroskedastic (Ehrmann et al.,

2011) to obtain information on the response of variables to each other.² The intuition behind it is that changes in the ratio of the variances of two structural shocks carry additional information and that each additional heteroscedastic regime adds more equation than unknowns to the model. In principle, two regimes are sufficient to identify the model.³

There are two conditions to be met in order for this methodology to solve the identification issue. First, the standard assumption in the VAR literature has to be valid, i.e. structural shocks have to be uncorrelated. Second, the contemporaneous impact matrix A has to be stable over time (i.e. stable across heteroscedastic regimes). Moreover, although the number of regimes identifies the system, this is only true up to a rotation of the matrix A . We therefore need to impose some additional restrictions to ensure that we pick the “correct” rotation, that is the rotation that represents the true underlying economic relationships among variables. In order to do so, in the literature some sign restrictions are usually imposed. Thus, following and Ehrmann et al. (2011), we assume that an increase in the two-year euro interest rate leads to an increase in the two-year US Treasury interest rate. This fairly uncontroversial restriction is imposed on the corresponding structural parameter contained in the A matrix, and will help us to correctly identify the structural parameters without restricting the coefficients of main interest. In fact, this sign restriction is not used to identify the matrix, but only to make sure to pick the solution that is economically meaningful and consistent with theory.

2.3 Identification of volatility regimes and estimation

Before the estimation, we need to determine the volatility regimes. In this paper we apply a statistical approach, where the determination of the regimes is data-driven. First, we estimate the reduced form model. Then we compute the rolling standard deviation for each reduced form residual (for each subsample) u_t . Then we calibrate the threshold for the rolling standard deviation that defines whether the residual should be classified into a high or low regime. The threshold, as well as the window used to compute the rolling standard deviation, are subsample-specific and they are reported in Table 1. For each subsample we identify five volatility regimes and we define

²Moreover, the specific form of heteroscedasticity is not of interest, as Rigobon (2003) shows that the estimates of the contemporaneous relationships are consistent regardless the form of heteroscedasticity.

³We need to impose at least a sign restriction to pick the right rotation of the A matrix. See later discussion.

them as follows: 1) all residuals are classified in low volatility (that is, the standard deviation of all three residuals is below the threshold); 2) only exchange rate equation residual is in high volatility, the other residuals are in low; 3) only the two-year euro rate equation residual is in high volatility, the other residuals are in low; 4) only two-year us Treasury interest equation residual is in high volatility, the other residuals are in low; 5) all residuals display high volatility. All observations that cannot be classified into any of the five regimes are excluded from the estimation.

To see whether our regime determination is supported by the data or not, we test formally for the constancy of the reduced form covariance matrix within each subsample. Following Lanne and Lütkepohl (2008) we perform a test on the joint null hypothesis that all five regimes have the same covariance matrix (Table 2). Moreover, we conduct pairwise likelihood ratio tests on the null hypothesis that any two regimes have the same variance-covariance matrix (Table 3). All null hypothesis are strongly rejected by the data but in one case.

After having determined the volatility regimes, we estimate the model as in Ehrmann et al. (2011), minimizing the following matrix norm:

$$\|g'g\| = \sqrt{tr(gg')} = \sqrt{vec(g)vec(g')},$$

$$\text{with } g = \sum_{k=1}^5 (A\Sigma_{u,k}A^{-1} - \Sigma_{\epsilon,k}), \quad (4)$$

where $\Sigma_{u,k}$ is the variance of the reduced form shocks in regime k , $\Sigma_{\epsilon,k}$ is the variance of the structural shocks in volatility regime k and A is the matrix of contemporaneous impact subject to the sign restriction defined in Subsection 2.2. We base statistical inference on bootstrap replications, drawing 150 times for each specific covariance matrices and for each draw estimating the coefficients by minimizing equation (6).

3 Results

Main results are contained in Table 4. Since the data are differenced, there is no persistence in the data. For this reason, we do not present impulse response functions, as they do not provide any additional insights.

3.1 Direct effects

Table 4 presents the estimated A-matrix for all four subsamples. These coefficients can be interpreted as the direct effects of the various shocks, thus not incorporating possible indirect effects via other asset prices. For ease of interpretation, we reverse the signs of the off-diagonal elements.

The results show that the relationship between the USD-EUR exchange rate and the US and the euro interest rates change over time. In tranquil times, the dynamic of the exchange rate is still dominated by the US, while in period of crisis, the exchange rate responds primarily to the shocks coming from the area where the crisis originates. Moreover, there are significant contemporaneous linkages across US and euro area interest rates. All the significant relations have the expected sign.

The following set of equations presents the result of contemporaneous response of the exchange rate to interest rate shocks for each subsample. They correspond to the estimates of the coefficients of the structural form model (2). We highlight parameters that are at least at the 90% significance level through bold font.

$$\begin{aligned}
 \textit{Pre-crisis period:} \quad e_t &= -0.23r_t^{EA} - \mathbf{0.23}r_t^{US} + \dots \\
 \textit{GFC period:} \quad e_t &= +0.24r_t^{EA} - \mathbf{0.29}r_t^{US} + \dots \\
 \textit{EC period:} \quad e_t &= +\mathbf{0.34}r_t^{EA} - \mathbf{0.23}r_t^{US} + \dots \\
 \textit{Post-OMT period:} \quad e_t &= +\mathbf{0.22}r_t^{EA} - \mathbf{0.27}r_t^{US} + \dots
 \end{aligned}$$

In the pre-crisis subsample, the US market dominates, as the coefficient attached to the Euro area is not significant. A positive shock to the 2-year Treasury rate leads to an appreciation of the US Dollar, that implies that the exchange rate decreases. With the unfolding of the global financial crisis, the importance of the US shocks in the determination of the exchange rate grows, while the euro area remains insignificant. Interestingly, there is no evidence of a sign reverse, that would indicate the presence of a “flight to quality” phenomenon. During the European sovereign crisis, the situation reverses and Euro area interest rates gain importance. The exchange rate now responds heavily to the development of the German interest rates, probably as a consequence of the fear of a possible Euro break-up. In fact, during the crisis, the flight to quality phenomenon was observed within Europe, from the peripheral countries to the core ones (Ehrmann and Fratzscher, 2017). Thus, an increase in the German interest rates was perceived as positive news from the financial markets, and this explains why it would lead to an appreciation of the Euro. Finally, after the announcement of the ECB’s

OMT program, the situation goes back to normality and the exchange rate dynamic is still dominated by the US market, although the European coefficient remains significant.

The following sets of equations presents the result of contemporaneous response of the two-year euro area and of the two-year US Treasury rates. The euro area rate respond to the exchange rate and to its US counterpart in the following way

$$\begin{aligned}
\text{Pre-crisis period:} \quad & r_t^{EA} = -0.2e_t - \mathbf{0.41}r_t^{US} + \dots \\
\text{GFC period:} \quad & r_t^{EA} = +0.24e_t - \mathbf{0.35}r_t^{US} + \dots \\
\text{EC period:} \quad & r_t^{EA} = -0.17e_t - 0.24r_t^{US} + \dots \\
\text{Post-OMT period:} \quad & r_t^{EA} = +0.08r_t^{EU} - \mathbf{0.31}r_t^{US} + \dots,
\end{aligned}$$

while the US Treasury rate responds to euro rate and exchange rate as follows:

$$\begin{aligned}
\text{Pre-crisis period:} \quad & r_t^{US} = 0.00e_t + \mathbf{0.34}r_t^{EA} + \dots \\
\text{GFC period:} \quad & r_t^{US} = 0.12e_t + \mathbf{0.27}r_t^{EA} + \dots \\
\text{EC period:} \quad & r_t^{US} = 0.20e_t + \mathbf{0.33}r_t^{EA} + \dots \\
\text{Post-OMT period:} \quad & r_t^{US} = 0.6e_t + \mathbf{0.21}r_t^{EA} + \dots .
\end{aligned}$$

The spillovers from the United States to the euro area are generally larger than in the other direction, consistent, among the others, with Chinn and Frankel (2005) and Ehrmann et al. (2011). The only exception being during the European crisis, when the coefficient on the US is not significant. Also here it is possible to notice that, in times of crisis, interest rates respond more heavily to shocks coming from the country where the crisis originates. In fact, moving from the global financial crisis to the European crisis period, the spillovers from the US to the euro become smaller (and, in this case, also insignificant), while when comparing the same two periods and taking into account the spillovers from the euro area to the US, we see that the former gains importance. Notably, both the spillovers from the US and the euro area decrease over time, pointing towards the direction of the two areas being less and less integrated.

3.2 Variance decomposition

To quantify the average economic significance of the different types of structural shocks and to understand which shock contribute the most in explaining the volatility of the endogenous variables, we compute the one-

week ahead forecast error variance decomposition (FEVD). As we have five different regimes, we obtain five forecast decomposition for each subsample. The FEVD shows how much of the forecast error variance of each variable can be explained by exogenous shocks to the other variables. Moreover, we compute a weighted average of the regime specific decomposition to measure the average importance of the shocks over each subsample, using the number of observations per regime as weights. Only this last set of results are reported in Table 5.

The weighted FEVD shows that the importance of interest rates in explaining the variation of the exchange rate gradually decreases, confirming the idea that the exchange rate was mainly dancing to its own tune (Ehrmann et al., 2014). In fact, in the pre-crisis period, the US and the Euro area shocks explain 16% of the variation of the exchange rate. This percentage is reduced to 7% during the global financial crisis and to 10% during the European crisis. Interestingly, also in the post-OMT period the explanatory power of both the interest rates do not go back to pre-crisis levels but remains lower. Concerning the two-year euro rate, in the pre-crisis period and during the European crisis 84% of its variation is explained by own shocks, while this percentage increases to 90% in the other two periods. Finally, regarding the two-year US interest rate, the variation explained by its shock is fairly stable, ranging from 87% to 92%.

3.3 The sensitivity of exchange rate and interest rates to macroeconomic news

In this section we present the effects of the exogenous control variables on the three endogenous ones. The values are obtained estimating the rows of the reduced form model (3) and using robust standard errors, in order to account for the heteroscedasticity present in the data. We report the effect of the news surprises on all three endogenous variables, for each of the subsamples. We do not interpret all single coefficients, as we mostly are interested in the general trend, and we only show significant coefficients in each subsample. In general, results show that, over time, all variable depict an increasing sensitivity to macroeconomic surprises coming from the euro area, while the responsiveness to US surprises is decreasing over time. This result point towards the idea that after the US nominal interest rates have reached the zero lower bound in December 2008 (i.e. during the global financial crisis subsample), the Fed's forward guidance policy and the large-scale asset purchases rounds have substantially constrained the behaviour of our endogenous variables, limiting their responsiveness (Swanson and Williams

2014a; Swanson and Williams 2014b).

The results for the USD-EUR exchange rate are reported in Table 7. On average, the exchange rate responds positively to output and inflation upward surprises in the Eurozone, that is, the Euro tends to appreciate in response to these surprises, and thus, the exchange rate tends to increase. Conversely, the Dollar tends to appreciate in response to US upward surprises in inflation and output, implying a decrease in the USD-EUR exchange rates. These findings are consistent with theory that predicts that surprises that imply an increase in domestic interest rates tend to appreciate the domestic currency while surprises that appreciate the foreign currency tend to depreciate it (Swanson and Williams, 2014b). The results support the finding described in Subsection 3.1: in the pre-crisis and in the global financial crisis subsamples the USD-EUR exchange rate does not respond to the euro area surprise releases. With the unfolding of the European crisis the situation changes and the exchange rate starts being responsive. This is consistent with the findings pointed out in section 3, i.e. that the euro area gains importance in determining the exchange rate only in the last two subsamples. Moreover, the exchange rate remains responsive also in the post-crisis period.

Concerning the surprises coming from the US, the exchange rate does not seem to be unusually responsive during the periods of crisis compared to the non-crisis subsamples. In the non-crisis period it appears that the economic developments in the US prove to play a larger role than the European counterpart in the determination of the exchange rate, in line with Ehrmann and Fratzscher (2005b). But with the unfolding of financial turmoils, the influence of macroeconomic surprises on the exchange rate does not change. This is again in line with Section 3: in period of crisis, the exchange rate continues to respond to US shocks in a similar manner as the non-crisis period. As already anticipated, one possible explanation is that the zero lower bound on nominal interest rates did not significantly constrain the behavior of the US Dollar despite being a significant constraint on short-term interest rates in the US. This finding is consistent with Swanson and Williams (2014b). All in all, we conclude that the exchange rate's responsiveness has increased with respect the euro area surprises, but that it was essentially unconstrained by the US zero lower bound.

The results for the two-year euro interest rate are reported in Table 8. The rate tends to increase in response to positive news about euro area's output and inflation and it tends to decrease in response to increase in unemployment, which, as pointed out by Swanson and Williams (2014b), is consistent with a Taylor-type reaction function for monetary policy. Moreover, the two year euro interest rate responds positively also to US news in

the same manner, with upward surprises in the overseas's output and inflation having a positive effect. From the results it appears that in the first subsample, the euro area reacts only to surprises components coming from the US, which is in line with Andersson et al. (2009)'s findings. Then, the sensitivity of the European rate to euro area surprises grows. Moreover, the results suggest that in period of crisis, the Euro rate responds primarily to the shocks coming from the area where the crisis originated. Furthermore, it is worth noting that, considering both sets of surprises together, the maximum sensitiveness of the euro interest rate is reached in the second subsample, i.e. during the global financial crisis. After that, the rate has started responding in a softer way to developments in the two areas, which can be seen as a first indication towards the rate as being constrained first by the US zero lower bound (the ZLB occurs in our second subsample, and the German yield starts being less responsive from the third subsample onwards), and then by the OMT announcement (and again, the OMT happens in the third subsample and in the post-OMT subsample the responsiveness to European developments drops).

The results for the USD-EUR exchange rate are reported in Table 9. Also the 2-year US Treasury interest rate tends to be positively related to upwards surprise on output and inflation and negatively related to unemployment surprises, coming from both the US and the euro area. Moreover, the US yields respond primarily to US developments. Interestingly, it is also possible to notice that, with respect to US surprises, the sensitivity of the Treasury has diminished over time, after reaching a pick during the global financial crisis subsample. This surprisingly responsiveness of the treasury during the financial crisis years is consistent with what documented by Swanson and Williams (2014a): it is only after 2011 (during our third subsample), i.e. after the Fed's forward guidance and the large-scale asset purchases rounds, that the sensitivity of the Treasury starts falling.

4 Conclusion

The aim of the paper is to study whether the relation between the US-Dollar/Euro exchange rate and US and euro area interest rates changes over time. We split the daily data sample starting in 2000 into four subsamples, following what the literature has identified as the main game-changers for the US and the euro area: a pre-crisis period, the global financial crisis, the European debt crisis, and a post-crisis sample. For each subsample we specify each asset price within a multifactor model and estimate the causal contem-

poraneous coefficients by exploiting the heteroskedasticity that is present in the daily data.

We find that the relationship between the exchange rate interest rates changes significantly across crisis and non-crisis times. In tranquil times, movements in the exchange rate are dominated by interest rate and macroeconomic new shocks from the US, while in periods of crisis, the exchange rate responds primarily to the shocks coming from the area where the crisis originates, implying an increased sensitivity to US developments during the global financial crisis and a higher responsiveness to euro area shocks during the European debt crisis. These findings contribute to the understanding of exchange rate behavior and indicate the existence of time-variation in the international transmission of macroeconomic shocks.

References

- Andersen, T. G., Bollerslev, T., Diebold, F. X., and Vega, C. (2003). Micro effects of macro announcements: Real-time price discovery in foreign exchange. *The American Economic Review*, 93(1):38–62.
- Andersson, M., Overby, L. J., and Sebestyén, S. (2009). Which news moves the euro area bond market? *German Economic Review*, 10(1):1–31.
- Baillie, R. T. and Bollerslev, T. (2000). The forward premium anomaly is not as bad as you think. *Journal of International Money and Finance*, 19(4):471–488.
- Bekaert, G., Hodrick, R. J., and Marshall, D. A. (1997). On biases in tests of the expectations hypothesis of the term structure of interest rates. *Journal of Financial Economics*, 44(3):309–348.
- Bekaert, G., Wei, M., and Xing, Y. (2007). Uncovered interest rate parity and the term structure. *Journal of International Money and Finance*, 26(6):1038–1069.
- Cecchetti, S. G. (2008). Monetary policy and the financial crisis of 2007–2008. *CEPR Policy Insight*, 21.
- Chinn, M. and Frankel, J. A. (2005). The euro area and world interest rates. Wisconsin and Harvard mimeo.
- Chinn, M. D. and Meredith, G. (2004). Monetary policy and long-horizon uncovered interest parity. *IMF Economic Review*, 51(3):409–430.

- Ehrmann, M. and Fratzscher, M. (2005a). Equal size, equal role? interest rate interdependence between the euro area and the united states. *The Economic Journal*, 115(506):928–948.
- Ehrmann, M. and Fratzscher, M. (2005b). Exchange rates and fundamentals: new evidence from real-time data. *Journal of International Money and Finance*, 24(2):317–341.
- Ehrmann, M. and Fratzscher, M. (2017). Euro area government bonds—fragmentation and contagion during the sovereign debt crisis. *Journal of International Money and Finance*, 70:26–44.
- Ehrmann, M., Fratzscher, M., and Rigobon, R. (2011). Stocks, bonds, money markets and exchange rates: Measuring international financial transmission. *Journal of Applied Econometrics*, 26(6):948–974.
- Ehrmann, M., Osbat, C., Stráský, J., and Uusküla, L. (2014). The euro exchange rate during the european sovereign debt crisis—dancing to its own tune? *Journal of International Money and Finance*, 49:319–339.
- Engel, C. (1996). The forward discount anomaly and the risk premium: A survey of recent evidence. *Journal of empirical finance*, 3(2):123–192.
- Engel, C. (2014). Exchange rate and interest parity. forthcoming in the handbook of international economics, vol. 4.
- Ferrari, M., Kearns, J., and Schrimpf, A. (2016). Monetary shocks at high-frequency and their changing fx transmission around the globe. *Unpublished working paper*.
- Fratzscher, M. (2009). What explains global exchange rate movements during the financial crisis? *Journal of International Money and Finance*, 28(8):1390–1407.
- Glick, R., Leduc, S., et al. (2013). The effects of unconventional and conventional us monetary policy on the dollar. Federal Reserve Bank of San Francisco working paper 2013-11.
- Glick, R., Leduc, S., et al. (2015). Unconventional monetary policy and the dollar: conventional signs, unconventional magnitudes. *Federal Reserve Bank of San Francisco Working Paper Series*, (2015-18).
- Kiley, M. T. (2013). Exchange rates, monetary policy statements, and uncovered interest parity: before and after the zero lower bound. Finance and Economics Discussion Series, 2013-17.

- Lanne, M. and Lütkepohl, H. (2008). Identifying monetary policy shocks via changes in volatility. *Journal of Money, Credit and Banking*, 40(6):1131–1149.
- Rey, H. (2015). Dilemma not trilemma: the global financial cycle and monetary policy independence. Technical report, National Bureau of Economic Research.
- Rigobon, R. (2003). Identification through heteroskedasticity. *The Review of Economics and Statistics*, 85(4):pp. 777–792.
- Rogers, J. H., Scotti, C., and Wright, J. H. (2014). Evaluating asset-market effects of unconventional monetary policy: a multi-country review. *Economic Policy*, 29(80):749–799.
- Rogers, J. H., Scotti, C., and Wright, J. H. (2016). Unconventional monetary policy and international risk premia. *Board of Governors of the Federal Reserve System (U.S.)*, International Finance Discussion Papers 1172.
- Sentana, E. and Fiorentini, G. (2001). Identification, estimation and testing of conditionally heteroskedastic factor models. *Journal of econometrics*, 102(2):143–164.
- Stavrakeva, V. and Tang, J. (2016). Exchange rates and the yield curve. *Mimeo*.
- Swanson, E. T. and Williams, J. C. (2014a). Measuring the effect of the zero lower bound on medium-and longer-term interest rates. *The American Economic Review*, 104(10):3154–3185.
- Swanson, E. T. and Williams, J. C. (2014b). Measuring the effect of the zero lower bound on yields and exchange rates in the uk and germany. *Journal of International Economics*, 92:S2–S21.
- Trichet, J.-C. (2010). State of the union: The financial crisis and the ecb’s response between 2007 and 2009. *JCMS: Journal of Common Market Studies*, 48(s1):7–19.

Appendix

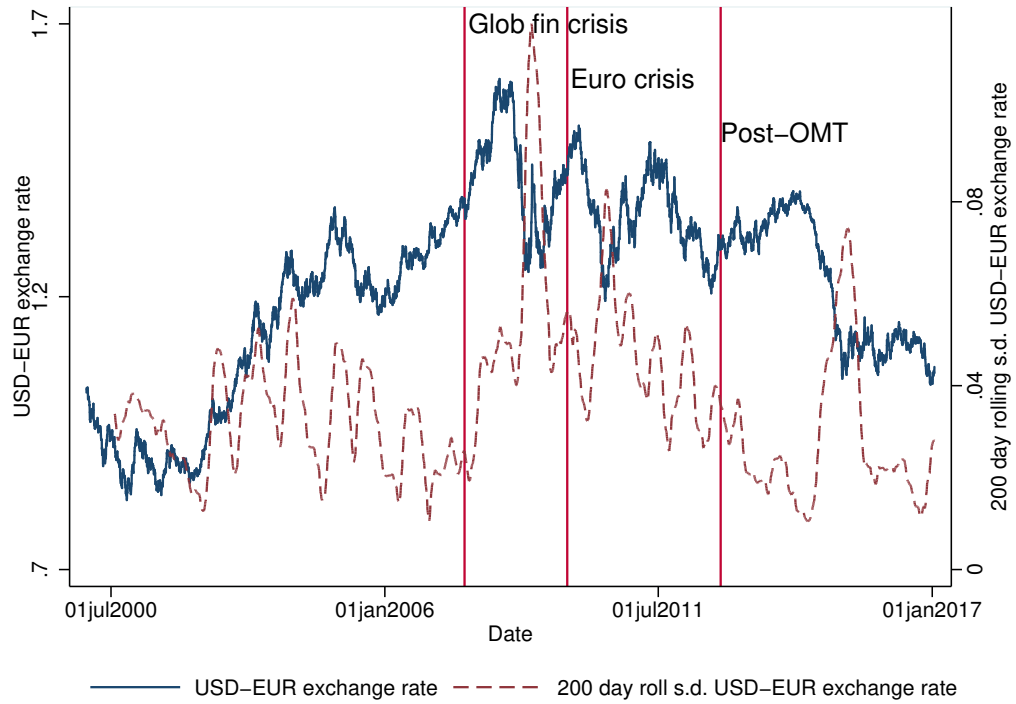


Figure 1: The figure shows the behaviour of the daily USD-EUR exchange rate (blue line, left axis) together with its 200 days rolling standard deviations (red dashed line, right axis). The sample period is 1 January 2000 until 3 November 2016.

Table 1: Window and threshold for the computation and the classification of reduced form residuals

	Pre-crisis	Global financial crisis	European crisis	Post-OMT
Window (weeks)	10	8	4	10
Threshold (stnd dev)	1.1	1	0.9	1

The table reports the window and the threshold needed to determine the volatility regimes used to identify the shocks in each subsample. First, we estimate the reduced form model. Then we compute the rolling standard deviation for each reduced form residual u_t , using the window reported above. Then we calibrate the threshold for the rolling standard deviation that defines whether the residual should be classified into a high or low regime. For each subsample we identify five volatility regimes and we define them as follows: 1) all residuals are classified in low volatility (that is, the standard deviation of all three residuals is below the threshold); 2) only exchange rate equation residual is in high volatility, the other residuals are in low; 3) only the two-year German interest rate equation residual is in high volatility, the other residuals are in low; 4) only two-year Treasury interest rate equation residual is in high volatility, the other residuals are in low; 5) all residuals display high volatility. All observations that cannot be classified into any of the five regimes are excluded from the estimation.

Table 2: Likelihood ratio tests for constancy of reduced form covariance matrix

	Pre-crisis	Global financial crisis	European crisis	Post-OMT
LR statistic	1069.5449	355.8850	558.8216	676.6048
p-value	(0.00)	(0.00)	(0.00)	(0.00)

The table shows results for all subsamples of likelihood ratio test on the null hypothesis that, within a subsample, all regimes have the same reduced form covariance matrix. The test statistic is obtained a follow:

$$LR = -2(\ln L_0 - \ln L_1)$$

where $\ln L_0$ is the restricted log likelihood function, that is the log likelihood function evaluated imposing the null hypothesis and $\ln L_1$ is the estimate of the unrestricted log likelihood function. P-values are in parentheses.

Table 3: Pairwise test for constancy of reduced form covariance matrix

		Regime			
Regime		1	2	3	4
Pre-crisis					
2	LR stat	55.53			
	p-value	(0.00)			
3	LR stat	13.19	19.11		
	p-value	(0.04)	(0.00)		
4	LR stat	43.22	41.82	7.62	
	p-value	(0.00)	(0.00)	0.27	
5	LR stat	142.37	51.47	18.83	24.68
	p-value	(0.00)	(0.00)	(0.00)	(0.00)
Global financial crisis					
2	LR stat	20.59			
	p-value	(0.00)			
3	LR stat	18.86	14.35		
	p-value	(0.04)	(0.03)		
4	LR stat	27.55	40.41	23.56	
	p-value	(0.00)	(0.00)	0.27	
5	LR stat	75.15	44.89	26.82	24.52
	p-value	(0.00)	(0.00)	(0.00)	(0.00)

Continued on next page

Table 3 – continued from previous page

		Regime			
Regime		1	2	3	4
European crisis					
2	LR stat	14.89			
	p-value	(0.02)			
3	LR stat	66.22	43.47		
	p-value	(0.04)	(0.00)		
4	LR stat	51.34	38.56	23.01	
	p-value	(0.00)	(0.00)	0.27	
5	LR stat	3132.37	70.30	24.97	35.40
	p-value	(0.00)	(0.00)	(0.00)	(0.00)
Post-OMT					
2	LR stat	62.6			
	p-value	(0.00)			
3	LR stat	46.35	68.99		
	p-value	(0.04)	(0.03)		
4	LR stat	36.63	35.98	81.85	
	p-value	(0.00)	(0.00)	0.27	
5	LR stat	104.62	29.61	76.08	49.35
	p-value	(0.00)	(0.00)	(0.00)	(0.00)

The table shows the results for all subsamples of the likelihood ratio tests on the null hypothesis that pairwise regimes have the same reduced form covariance matrix. The first row from the top and the first column from the left indicate the regime. For instance, the cell at the intersection between regime 1 and regime 2 (the value shown in that cell is 14.89) contains the likelihood ratio test statistic of the test on the null hypothesis that regime 1 and regime 2 have the same variance-covariance matrix. P-values are in parentheses.

Table 4: Contemporaneous effects among endogenous variables

Response	Impulse		
	USD-EUR exchange rate	Two-year euro area rate	Two-year US interest rate
Pre-crisis			
USD-EUR ex rate	1.00	-0.23	-0.23
p-value	.	.	.b
Two-year euro rate	-0.02	1.00	0.41
p-value	.	.	.a
Two-year US rate	0.00	0.34	1.00
p-value	.	.a	.
Global financial crisis			
USD-EUR ex rate	1.00	0.24	-0.29
p-value	.	.	.b
Two-year euro rate	0.17	1.00	0.35
p-value	.	.	.a
Two-year US rate	0.12	0.27	1.00
p-value	.	.b	.
European crisis			
USD-EUR ex rate	1.00	0.34	-0.23
p-value	.	.c	.b
Two-year euro rate	0.24	1.00	0.24
p-value	.	.	.
Two-year US rate	0.20	0.33	1.00
p-value	.	.c	.
Post-OMT			
USD-EUR ex rate	1.00	0.22	-0.27
p-value	.	.b	.c
Two-year euro rate	0.08	1.00	0.31
p-value	.	.	.a
Two-year US rate	0.06	0.21	1.00
p-value	.	.a	.

The table shows the estimated impact effects of structural shocks on the endogenous variables, for all four subsamples based on a structural VAR identified through heteroskedasticity. Impulse variables are in columns, response variables are in rows. For ease of interpretation, the signs of the off-diagonal elements are reversed. The sample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 3 Nov 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

Table 5: Weighted forecast error variance decomposition

	USD-EUR exchange rate	Two-year euro area rate	Two-year US interest rate
Pre-crisis			
USD-EUR ex rate	0.84	0.08	0.09
Two-year euro rate	0.00	0.84	0.15
Two-year US rate	0.00	0.10	0.90
Global financial crisis			
USD-EUR ex rate	0.93	0.02	0.05
Two-year euro rate	0.01	0.90	0.09
Two-year US rate	0.01	0.07	0.92
European crisis			
USD-EUR ex rate	0.90	0.06	0.04
Two-year euro rate	0.07	0.84	0.09
Two-year US rate	0.05	0.08	0.87
Post-OMT			
USD-EUR ex rate	0.91	0.04	0.05
Two-year euro rate	0.01	0.90	0.09
Two-year US rate	0.01	0.07	0.92

The table shows the weighted forecast error variance decompositions over an horizon of one week for each subsample. For each subsample the weighted FVED is computed averaging over the forecast error variance decompositions calculated for each regime, using the number of observations per regime as weights, based on a structural VAR identified through heteroskedasticity. In each subsample, the classification of the regimes is as follows: in regime 1 all structural shocks have low volatility; in regime 2 the USD-EUR exchange rate shock is in high volatility, the other two shocks are in low volatility; in regime 3 the 2-year euro area interest rate is in high volatility, the other two shocks are in low volatility; in regime 4 the 2-year Treasury interest rate is in high volatility, the other two shocks are in low volatility; in regime 5 all shocks are in high volatility. The threshold that defines the high volatility is subsample-specific (see Table 1).

Table 6: List of included data releases

<i>Euro area surprises</i>	<i>United States surprises</i>
European Commission Business Climate Indicator	CPI Urban Consumers YoY NSA
ECB Current Account Net WDA SA	CPI Urban Consumers Less Food and Energy YoY NSA
European Commission Consumer Confidence Indicator Eurozone	Personal Consumption Expenditure CPI YoY SA
Euro Consumer Price Index YoY	CPI Urban Consumers MoM SA
Euro Consumer Price Index MoM	of Michigan Consumer Confidence Indicator
Eurozone BOP Current Account Net NSA	US Government Budget Balance (FED)
Eurostat Eurozone Govt Debt as a % of GDP	American Consumer Spending Growth Rates MoM SA
Eurostat New Orders Eurozone (Manufacturing Industries) YoY	Markit Manufacturing PMI SA
Gross Fixed Capital Formation (Chained 2010 Prices) QoQ	Markit Composite PMI SA
European Commission Services Confidence Indicator	Markit Services PMI Business Activity SA
Markit Eurozone Composite PMI SA	Industrial Production MoM 2007=100 SA
Markit Eurozone Manufacturing PMI SA	Core Producer Price Index
Markit Eurozone Services PMI SA	Producer Price Index - Finished Goods
Eurostat Retail Sales Volume YoY WDA	Initial Jobless Claims SA
Eurostat Retail Sales Volume MoM SA	Housing Starts/Permits
ZEW Eurozone Expectation of Economic Growth (Econ. Sentiment)	GDP Chained 2009 Dollars QoQ SAAR
Eurozone Trade Balance with non Eurozone Countries	PPI Finished Goods SA MoM%
ECB Money Supply M3 YoY	PPI Final Demand MoM SA
ECB M3 Money Supply 3 Month Moving Avg	Capacity Utilization % of Total Capacity
PPI Eurozone Industry Ex Construction YoY	Business Inventories MoM SA
PPI Eurozone Industry Ex Construction MoM	Avg. Hourly Earnings YoY% SA
Eurostat Unemployment Rate	Avg. Hourly Earnings MoM% SA
GDP SA QoQ (real 2010 prices SA)	Construction Spending Total SA
	Construction Spending Total MoM SA
	Durable Goods New Orders Industries MoM SA
	Conference Board Leading Indicators MoM
	Productivity Output Per Hour Nonfarm Business Sector QoQ SA
	Unit Labor Costs Nonfarm Business Sector QoQ % SAAR
	Retail Sales (Less Auto and Gas Stations) SA MoM % Change
	Personal Income MoM SA
	Nonfarm Payrolls Total MoM SA
	Difference Between Exports and Imports

Table 7: Effects of selected surprises on the USD-EUR exchange rate

	Pre-Crisis	Global financial crisis	European crisis	Post-OMT
EC Business Climate Ind	0.93	1.78*	0.91	2.18**
EC Consumer Confidence Ind	0.00	-0.00	0.00*	0.00**
Euro CPI YoY	0.50	2.08	2.54	-4.13*
EC Services Confidence Ind	-0.00	0.00	-0.00 ^b *	-0.00 ^b **
Retail Sales Volume YoY	-0.00	0.00	0.00**	-0.00*
Retail Sales Volume MoM	-0.00	0.09	-0.16	-0.00***
Trade Balance with non Eurozone	-0.00	0.00*	-0.00*	0.00
M3 Money Supply 3Month Mov Avg	0.51	-0.63	3.93***	-1.11
Unemployment Rate	-0.65	-1.76	-5.17	-3.48**
<i>United States surprises</i>				
CPI Urban Consumers YoY	0.30***	-0.09	0.09	0.27
Personal Cons Expenditure CPI YoY	0.10	-0.29	0.26	-0.44*
US Government Budget Balance	0.10*	0.01	-0.12	0.09
Exports/Imports difference	-0.35***	-0.22	0.22	-0.10
PPI-Finished Goods	0.04	0.63	0.89*	-0.20
PPI Finished Goods MoM	0.06	-0.55	-1.06**	0.30
Business Inventories MoM	-0.04	0.24	-0.11	-0.22*
Construction Spending Tot MoM	0.00	0.38*	0.15	0.06
Personal Income MoM	-0.27*	0.02	-0.00	0.01
Nonfarm Payrolls Total MoM	0.07	-0.32	0.20	-0.86**

The table shows the effects of statistically significant variables in each subsample on the USD-EUR exchange rate from the baseline VAR, obtained from estimating the first row of the reduced form model (3) with robust standard error, in order to account for heteroscedasticity:

$$e_t = c + \alpha e_{t-1} + \beta r_{t-1}^{EA} + \gamma r_{t-1}^{US} + \delta x_t + u_{e,t},$$

where $e = \log \Delta$ USD-EUR exchange rate; $r^{EA} = \Delta$ 2-year German yield and $us = \Delta$ 2-year Treasury yield. Only estimates of vector δ , i.e. of the exogenous regressors have been reported here. All surprises have been standardized. ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

Table 8: Effects of selected surprises on the two-year German rate

	Pre-Crisis	Global financial crisis	European crisis	Post-OMT
EC Business Climate Ind	-0.80	3.76 ^a	0.50	1.05
EC Consumer Confidence Ind	0.00	-0.00 ^{***}	0.00 ^{***}	-0.00 ^{**}
Euro CPI YoY	0.41	10.70 ^{**}	3.22	-1.08
Gross Fixed Capital Formation QoQ	0.32	-3.37	1.13	2.47 ^{**}
EC Services Confidence Ind	-0.00	0.00 ^{***}	-0.00 ^{**}	0.00
Retail Sales Volume YoY	0.00	0.00	0.00 ^{**}	0.00
Retail Sales Volume MoM	-0.00	-0.94	-0.53 [*]	-0.00 ^{***}
M3 Money Supply 3 Month Mov Avg	-0.06	1.16	2.28 ^{**}	-1.92
PPI Industry Ex Construction YoY	-0.09	1.23	-2.20	-2.77
Unemployment Rate	0.25	1.57	-5.41 [*]	-1.84
GDP SA QoQ	0.28	3.71 [*]	-0.59	0.84
<i>United States surprises</i>				
CPI Urban Consumers YoY	-0.21	0.62 ^{**}	-0.35	-0.08
CPI Urban Cons, no Food&Energy	-0.01	0.09	-0.02	-0.08
Personal Cons Exp CPI YoY	0.12	-0.14	0.30	-0.03
CPI Urban Consumers MoM	0.12	-0.76 ^{**}	0.11	0.06
US Con Spending Growth Rates MoM	-0.01	0.51 [*]	0.10	-0.01
Trade Balance of Goods and Services	0.22 ^{**}	0.09	0.51 ^{**}	-0.03
Core PPI	0.30 ^{***}	0.53	-0.01	0.29
Initial Jobless Claims	-0.16 ^{***}	-0.24 ^{***}	-0.20 [*]	0.03
GDP QoQ	-0.37	0.72 ^{***}	0.02	0.25
Capacity Utilization	0.09	0.61 [*]	0.25	-0.14
Construction Spending Total MoM	-0.15	0.22 ^c	0.06	0.20
Durable Goods New Orders MoM	-0.01	-0.07	-0.14	0.31 [*]
Prod Outpout Per Hour Nonfarm QoQ	-0.09	0.26	0.73 [*]	0.02

The table shows the effects of statistically significant variables in each subsample on the 2-year German yield from the baseline VAR, obtained from estimating the second row of the reduced form model (3) with robust standard error, in order to account for heteroscedasticity:

$$r_t^{EA} = c + \alpha e_{t-1} + \beta r_{t-1}^{EA} + \gamma r_{t-1}^{US} + \delta x_t + u_{EA,t}$$

. where $e = \log \Delta$ USD-EUR exchange rate; $r^{EA} = \Delta$ 2-year German yield and $r^{US} = \Delta$ 2-year Treasury yield. Only estimates of vector δ , i.e. of the exogenous regressors have been reported here. All surprises have been standardized. .^{***}, .^{**} and .^{*} denote significance at the 1%, 5% and 10% levels, respectively.

Table 9: Effects of selected surprises on the two-year US Treasury rate

	Pre-Crisis	Global financial crisis	European crisis	Post-OMT
EC Business Climate Ind	-1.35	2.46**	-1.69**	-0.55
EC Consumer Confidence Ind	0.00	-0.00	0.00	-0.00***
CPI YoY	-2.51	9.21	5.97	9.71**
CPI MoM	1.49	-2.07	0.10	-2.75
Retail Sales Volume YoY	0.00	0.00	0.00**	-0.00
Retail Sales Volume MoM	-0.00**	0.18	-1.02*	0.00***
Trade Balance with non Eurozone	0.00**	0.00	0.00	0.00
PPI Industry Ex Construction YoY	0.19	-1.66	-6.11*	3.60**
PPI Industry Ex Construction MoM	-0.64	1.46	8.02***	-3.09*
<i>United States surprises</i>				
CPI Urban Consumers YoY	0.08	1.15*	-0.53	-0.66
CPI Urban Cons, no Food&Energy	0.25v	0.45	-0.10	-0.21
CPI Urban Consumers MoM	-0.03	-1.26v	0.15	0.71*
Uni of Michigan Cons Confidence Ind	0.06	0.13	-0.40*	-0.01
Con Spending Growth Rates MoM	-0.13	1.05***	-0.03	-0.08
Trade Balance of Goods&Services	0.11	0.16**	0.26	-0.28**
Core PPI	0.26***	0.04	0.03	0.09
PPI - Finished Goods	-0.17	-0.72**	-0.58	2.57
Initial Jobless Claims	-0.27***	-0.14*	-0.33***	-0.05
Housing Starts/Permits	-0.08	-0.07	0.43	0.32**
PPI Finished Goods	-0.11	0.80**	0.93**	-2.57
Business Inventories MoM	-0.32*	-0.13	0.20	0.27*
Construction Spending Tot MoM	0.03	0.49**	-0.06	0.10
Prod Outpout Per Hour Nonfarm QoQ	-0.07	0.13	0.63*	0.29
Retail Sales (no Auto&Gas Stations)	0.00	-0.10	-0.04	0.55***
Personal Income MoM	0.27**	-0.53***	-0.11	-0.01
Nonfarm Payrolls Total MoM	0.46***	0.44	0.20	1.35***

The table shows the effects of statistically significant variables in each subsample on the 2-year Treasury yield from the baseline VAR, obtained from estimating the third row of the reduced form model (3) with robust standard error, in order to account for heteroscedasticity:

$$r_t^{US} = c + \alpha e_{t-1} + \beta r_{t-1}^{EA} + \gamma r_{t-1}^{US} + \delta x_t + u_{US,t},$$

where $e = \log \Delta$ USD-EUR exchange rate; $r^{EA} = \Delta$ 2-year German yield and $r^{US} = \Delta$ 2-year Treasury yield. Only estimates of vector δ , i.e. of the exogenous regressors have been reported here. All surprises have been standardized. ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.