Assessing Euro Crises from a Time Varying International CAPM Approach*

Richard T. Baillie[a,b,c,†] Dooyeon Cho[d,‡]

[a]Department of Economics, Michigan State University, USA

[b]School of Economics and Finance, Queen Mary University of London, UK

[c]Rimini Center for Economic Analysis, Italy

[d]Department of Economics, Sungkyunkwan University, Seoul, Republic of Korea

Abstract

This paper initially reviews the current empirical literature on the Euro exchange rate. We consider the relationship between the euro and other floating currencies in terms of excess returns on bond markets and also the relationship between the euro-dollar and the US and European equity markets. One novelty in the paper is to consider the variation in the euro-dollar rate from an international capital asset pricing model (CAPM) perspective. The second new innovation is to use a kernel weighted time varying parameter regression approach which allows structural parameters and risk premium terms to evolve over time. We find evidence that the euro-dollar rate is substantially influenced by equity markets in the US and in the Eurozone.

JEL Classification: C22; F31; G01; G15

Keywords: Euro-US dollar exchange rate; International CAPM with factors; kernel weighted time-varying parameter regression

*The authors are very grateful for helpful comments from two anonymous referees, and from the editor Mark Jensen. Any remaining errors are our responsibility.

†Department of Economics, Michigan State University, East Lansing, MI 48824-1024, USA; Tel: +1 517 355 1864, Fax: +1 517 432 1068, E-mail: baillie@msu.edu.

‡Corresponding author. Department of Economics, Sungkyunkwan University, Seoul 03063, Republic of Korea; Tel: +82 2 760 0148, Fax: +82 2 760 0946, E-mail: dooyeoncho@skku.edu.

1 Introduction

Despite the existence of the Euro since 1999 and the great policy issues surrounding its performance and behavior, there have been surprisingly few econometric studies of the euro exchange rate. However almost sixteen years of floating exchange rate data now exists and allows some of the basic long run theories of exchange rate determination to be assessed and tested. This seems especially relevant given the discussion in de Grauwe (2016), which provides substantial background information on the history and political economy of the euro exchange rate. As noted by de Grauwe (2016), there have been several periods of turbulence in the Eurozone that has led to substantive concerns that the European Monetary Union (EMU) would not survive. As is well known in the study of international finance, the history of fixed exchange rates and monetary unions does not provide much optimism for the long term survival of the EMU. This is in the context that the current EMU has existed sixteen years, while the Bretton Woods system, which was one of the longest lasting managed regimes, lasted from 1944 through the early 1973. However, EU officials and policy makers have frequently in recent times appeared sanguine about the status quo of the euro and have stated that the worst of the euro crisis has now ended. They have appeared to take comfort by the ostensibly lower levels of turbulence and volatility in the euro exchange rate and associated bond markets since 2014. This has been attributed to institutional changes of tighter fiscal policies and the reduction of imbalances within the banking sector since the sovereign debt crisis of 2008. The article by de Grauwe (2016) provides a less optimistic view and considers the distinct possibility that the occurrence of the sovereign debt crisis combined with poorly designed fiscal policies are primarily responsible for the continuing low growth performance of the Eurozone. Furthermore the levels of accumulated debt may well herald a new era of possible crisis for the euro rate.

This current study to some extent complements the mainly fiscal explanations provided by de Grauwe (2016), and we investigate some alternative asset pricing explanations of instabilities and crises of the euro-dollar rate. We focus on some of the asset market linkages and the interactions between equity, bond and currency markets. One of the motivations for this analysis is in terms of an international $CAPM$, or $I-CAPM$ framework where the excess returns on the Eurozone equity market is related to the excess returns on the market as described by the US equity returns. In keeping with the $CAPM$ we include other various factors which include volatility of returns in both US and European equity markets. One novelty of the work is that both daily and monthly econometric analysis is performed and the kernel weighted regression methods with time varying parameters turn out to be particularly useful in identifying periods of instabilities and crisis in the Eurozone. Some of the results suggest subsequent macro factor analysis to identify and investigate the causes of crises and instabilities.

The plan of the rest of the paper is as follows: the next section summarizes some of the existing empirical work on the euro and the standard monetary model and purchasing power parity theories. Section 3 examines some of the empirical regularities and features of the excess euro - dollar returns vis-à-vis the bond markets and also vis-à-vis the inflation differentials. The use of the time varying parameter model is forcibly illustrated at this point since there is clearly considerable parameter variation over time. Some of the variation is directly attributable to the financial market crisis events. Then section 4 of the paper develops the idea of the possible use of the $I-CAPM$ in the context of modeling the euro-dollar exchange rate. These models are again estimated from a time varying parameter approach for both daily and monthly data. There appears to be some relatively strong interactions between both the US and European stock markets on the euro. These interactions became particularly strong in 2009 around the financial crisis and lasted for about four years. Section 5 summarizes the findings and offers brief suggestions for future research.

2 Previous Empirical Studies of the Euro

As previously mentioned, surprisingly few econometric studies of the euro exchange rate have appeared in the literature. Undoubtedly a primary reason for this has been the lack of substantial data and we briefly summarize some of the relatively few other empirical studies in this area. Several articles have taken a bits and pieces approach to analyzing component features of the basic model for the euro. In particular, it is possible to use data before the introduction of the euro when the European Monetary System (EMS) existed between 1979 and 1993 and was based upon a policy of having target rates against the Deutschmark, with relatively frequent realignments and devaluations of other currencies when deemed necessary. These policies of trying to maintain a band of the floating exchange rates were designed to achieve convergence to the single currency, which was eventually achieved in January 1999. This occurred despite occasional severe crises, such as the one in September 1992 when the UK pound and Italian lira were forced to leave the EMS and again in September 1993 when the French franc suffered a sustained depreciation and the EMS bands were widened to plus or minus 15% to accommodate extreme volatility within the European currency market. An obvious area of interest is to construct an artificial, or synthetic euro series before 1999 to join with actual data since 1999, and to hence form longer data series. Lopez and Papell (2007) use such an approach to generate data before 1999 when the informal EMS was in operation and considered the convergence towards purchasing power parity (PPP) within the Eurozone and between the Eurozone and its main partners using panel data methods. They generally rejected the hypothesis of a unit root in the Eurozone real exchange rates, and therefore find evidence of long run convergence towards PPP. This was found to hold for different numéraire currencies, as well as in the Eurozone plus the United States with the dollar as the numéraire currency, starting between 1996 and 1999. Hence these authors found evidence that the process of convergence towards PPP began as early as immediately after the currency crises of 1992 and 1993, and the adoption of the Maastricht Treaty concerning the single market of the Eurozone.

A related article by Koedijk, Tims and van Dijk (2004) studied the impact of the introduction of the euro in 1999 on the behavior of real exchange rates. Over the period 1973 through 2003, the authors exploited the cross sectional dependence across real exchange rates and allowed for heterogeneity in the rates of mean reversion. They find evidence in favor of PPP for the full panel of real exchange rates. They also examine the artificial or "synthetic" euro against several other major currencies over the period 1979–2003 and again found support for the PPP hypothesis for the full panel of real exchange rates.

Chinn and Moore (2011) consider an order flow approach and use the conventional specification of a model with monetary fundamentals and the Evans and Lyons (2002) microstructure approach, which is augmented with order flow variables. The data are inter-dealer order flow on the US dollar-euro and dollar-yen. They find the augmented or "hybrid" model compares favorably with the random walk specifications, and has some predictable content.

Chen, Fausten and Wong (2011) analyze possible cointegration between the euro-dollar exchange rate and consider fundamentals from the monetary model of exchange rate determination. They find that both the short-run variables associated with price stickiness and long-run fundamentals associated with secular trends tend to affect the exchange rate path. An alternative approach is by Bianco, Camacho and Quiros (2012) who use a fundamentals based econometric model for the weekly changes in the euro-dollar rate and mixing economic variables quoted at different frequencies. These authors also found that their model improves on simple random walk formulation for out of sample prediction.

A further paper by Beckmann, Belke and Kuhl (2011) investigates the temporal stability of the relationship between the Deutschmark-US dollar exchange rate and macroeconomic fundamentals. They conclude there is an absence of stable long-run equilibrium relationship among fundamentals and exchange rates, since the breakdown of Bretton Woods. They also find several regime breaks, and some related findings are reported by Canarella, Miller and Pollard (2014). Another fundamental type analysis is provided by Sartore et al (2002), who consider the specification and estimation of an econometric model of the US dollar-Euro real exchange rate in $VECM$ form. The authors endogenize the long term interest rate differential between US and Euro GDP annual growth rate in addition to the exchange rate. In this way the foreign exchange market, the money market and the goods market, are modelled jointly. The authors estimate a $VECM$ from monthly data from 1990 through 1999, which also includes the former EMS, or "snake" in the estimation period. The models include a structural break in September 1992 and there is some doubt as to whether there are really three cointegrating relationships in the model which includes the synthetic artificially generated euro series beginning in 1990.

There are several other studies considering the euro and non standard fundamentals; see Camarero and Ordonez (2012), who analyze the influence of productivity differentials on the real dollar–euro exchange rate. They use $ESTAR$ models during the period 1970–2009 and find that the dollar-euro real exchange rate shows nonlinear mean reversion towards the fundamentals represented by the productivity differential.

3 The Euro and Bond Markets

As is well known, there has been a general lack of success of macro fundamental models of exchange rate determination. Many of the points concerning the macroeconomic aspects of the euro are to be found in de Grauwe (2016). For this reason we concentrate in this paper on the linkages between asset markets to throw light on the behavior of the euro exchange rate. There are many aspects to the asset market behavior of the euro currency which are necessary to consider in the formulation of an international $CAPM$ approach. We first consider the euro in relation to bond markets in different countries or regions. In the following we define s_{t+1} as the natural logarithm of the spot exchange rate at time $t+1$ which is the number of units of foreign currency per one Euro, while i_t is the domestic nominal interest rate, and an asterisk denotes a foreign equivalent. Finally, E_t is the conditional expectation based on a sigma field of information available at time t.

While uncovered interest rate parity is a fundamental relationship for floating exchange rate regimes, there appears to be no existing detailed analysis of this for the euro exchange rate. Furthermore, many studies such as Baillie and Kiliç (2006), Lothian and Wu (2011), Baillie and Cho (2014), etc have shown that there are strong asymmetries occurring in the different regimes which are often associated with the sign of the interest rate differentials. These appear to affect deviations from uncovered interest rate parity and possibly time varying risk premium in non linear ways. It has been standard to test the theory from the regression equation $\Delta s_{t+1} = \alpha + \beta(i_t^* - i_t) + u_{t+1}$, where the theory of UIP implies $\alpha = 0$, $\beta = 1$ and u_{t+1} being serially uncorrelated. The article by Baillie and Cho (2014) considered estimation of the time varying parameter (TVP) forward premium regression of

$$(s_{t+k} - s_t) = \alpha_t + \beta_t(i_t^* - i_t) + u_{t+k}, \tag{1}$$

for $k = 1$, so that the sampling interval exactly matches the time period of the forward contract. One of the highlights of the Baillie and Cho (2014) study was the finding of temporal variation in the estimated β_t slope coefficient for currencies vis-à-vis the US dollar. Baillie and Cho (2014) did not consider the euro in their study, and this article extends their analysis to the case of the euro exchange rate against the currencies of Australia, Canada, Denmark, Japan, New Zealand, Norway, Switzerland, UK and the US with the euro as the numéraire currency. The data is monthly for spot exchange rates and one month forward exchange rates, and is obtained from *Datastream*. The estimation period is from January 1999, which is the first month of the formation of the euro, through September 2015, which provides a total of $T = 201$ monthly observations. A full description of the econometric technique is to be found in Baillie and Cho (2014) and their approach utilized and moderately extended some recent results on time varying parameter (TVP) autoregressions due to Giraitis, Kapetanios and Yates (2014, 2015). The methodology for this paper is quite similar except for calculation of covariance matrices adjusting for moving average and heteroskedastic error processes. These issues are relevant for some of the daily data which are used in this paper. A very brief summary of some of the new results is provided in the appendix of this paper.

In order to appreciate the properties of the various $\hat{\beta}_t$ for each currency, figure 1 shows the time plots of the estimated slope coefficient, while a summary of the number of the statistically significant violations of UIP, above and below the generated 95% confidence intervals for each time period, are summarized in table 1. Following the results of Giraitis, Kapetanios and Yates (2014) and Baillie and Cho (2014), the bandwidth was chosen as the nearest integer to $H = T^{1/2}$, so that $H = 14$ for the monthly data in this study. As seen from figure 1, there appears to have generally been substantial variation in the estimated β_t over time. Interestingly, table 1 indicates that there have been very few statistically significant departures from UIP for the currencies of Australia, Japan, Switzerland and the UK. At a very simple level this suggests that the UK pound is in some sense is “closer” to the euro in basic parity terms than commonly thought by many financial commentators. On the other hand Denmark has the classic forward premium paradox vis-à-vis the euro with home country bias for every time period. Conversely, New Zealand and Norway (which has been a high interest target currency in carry trades) have 22% and 15% of their period with statistically significant excess returns over UIP respectively.

The relatively smooth values of $\hat{\beta}_t$ obtained by the kernel weighted regression show slow changes that suggest some degree of predictability of the excess returns. Hence a possibility that was investigated was that there is macroeconomic information that allows prediction of the $\hat{\beta}_t$ through regression on a set of information as in $\hat{\beta}_t = \mathbf{x}_t'\boldsymbol{\theta}_t$. While the $\hat{\beta}_t$ are quite strongly autocorrelated, a regression of the $\hat{\beta}_t$ on the contemporaneous and lagged values of US and Eurozone inflation did not produce significant results. rates was not significant. Inflation for the US was defined as the differenced logged consumer price index (CPI) series as usual. For the Eurozone inflation is defined as the harmonized indices of consumer prices ($HICP$) in the Eurostat website. The results are not reported in this paper but are available from the authors on request.

Since the US is the other country identified in the upcoming $I-CAPM$ section of this paper, it was decided to focus on some analysis of daily returns data. Hence figure 2 documents the time paths of daily $\hat{\beta}_t$ derived from the model for daily euro-dollar excess returns from UIP

given by the model in equation (1), where now $k = 21$ to allow for 21 working days in each month and in the forward contract. In this case the autocorrelation structure of the error term is replaced with $E(u_t u_{t+j}) = 0$, for $j > k$. This restriction is used to provide appropriate consistent covariance matrix of the estimated TVP regression coefficients, which is described in the appendix. The sample size is $T = 4,368$ and the bandwidth parameter is $H = 66$.

The estimated β_t parameter for the excess returns over UIP for the euro-dollar becomes excessively volatile after 2010 with strong reversals of UIP. Figure 2 graphically illustrates the increasing degree of uncertainty and high variability of $\hat{\beta}_t$ towards the end of the sample period. This confirms the findings in table 1 that the behavior of $\hat{\beta}_t$ is quite different for the US dollar compared with the other currencies. It appears that uncertainty following the financial crisis is especially relevant for the euro against the US dollar.

The next possibility considered was that the euro-dollar rate was directly influenced by the performance of the US and European equity markets through direct capital flows for equity investments. In the following, $ESTOXX$ denotes a variable measuring the returns on the domestic (European markets) and $S\&P$ denotes returns on the US equity market.[1] The data is both monthly and daily for the $S\&P500$ index and $EURO\ STOXX\ 50$ index, and is also obtained from *Datastream*. Hence we estimate the TVP kernel weighted regression of

$$(s_{t+1} - s_t) = \alpha_t + \beta_t(i_t^* - i_t) + \gamma_t S\&P_t + \delta_t ESTOXX_t + \theta_t \Delta VIX_t + u_{t+1}, \qquad (2)$$

from monthly data and the results are reported in table 2 and figure 3. It is interesting to note that the time plot of the estimated β_t does not fundamentally change very much following the introduction of the other two equity market return variables. However, the number of periods where UIP is rejected are approximately 25% of the time in equation (1) and only 7% of the observations in equation (2). The probable reason is that the presence of the two equity return series reduces the excess returns over UIP and could be proxying some form of time dependent risk premium on the euro-dollar rate. However it should also be noted that the variability of the estimated β_t again increases considerably from 2012 onwards.

Most theories of time dependent risk premium generally use some measures of volatility. For

[1]The $EURO\ STOXX$ 50 Index is a Blue-chip index for the Eurozone and covers 50 stocks from the 12 Eurozone countries of Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain.

monthly data analysis we use the standard VIX measure of overall market volatility. While for daily data analysis we use the conditional variance from a $FIGARCH$ model of Baillie, Bollerslev and Mikkelsen (1996), for generating measurable volatility on the euro-\$ exchange rate, which are reported in table 3. Details of the estimated coefficients in equation (2) are provided graphically in figures 3 and 4, and also in table 2. The general results for $\hat{\beta}_t$ do not apparently change with the introduction of a further explanatory variable measuring the overall financial market volatility, the VIX. In particular, $\hat{\gamma}_t$ represents the effect on the euro-dollar rate of a change in returns on the $US\ S\&P500$ equity market. The parameter $\hat{\gamma}_t$ is generally positive and significantly different to zero coefficient, which implies that an appreciation of the US stock market is related to a euro depreciation vis-à-vis the US dollar. This would be consistent with generally high returns on the $S\&P500$ leading to an increased demand for US \$ to invest and funds flow out of the Eurozone to buy \$ denominated assets. Similarly the path of the estimated $\hat{\delta}_t$ indicates that increased returns on $EUROSTOXX$ in the Eurozone lead to an appreciation of the euro, which is consistent with increased demand for euros being motivated by an increase demand to purchase Eurozone equity assets.

It is important to acknowledge that the equation (2) is essentially a reduced form equation. Also, since the returns are all close to being serially uncorrelated, any VAR analysis is not very helpful. A corresponding equation linking equity market returns to the euro rate is from the $I-CAPM$ in the next section.

4 The Euro in an International CAPM Formulation

As previously discussed, this paper concentrates on the linkages between asset markets and the euro. It should be noted that using the concept of $I-CAPM$ is not a new idea and has been around in the exchange rate literature since Frankel (1982), Frankel and Engel (1984) and Engel and Rodrigues (1989). These studies typically allowed the ex post rates of return to vary freely, but restricted the portfolio to only consist of different currencies and for the supply of each asset to be measured as cumulated government debt. As noted by Frankel (1982) and Engel and Rodrigues (1989), this approach is quite limited due to data availability issues. Our framework is considerably more general in that we take into account portfolios to include bonds and equity returns. The study by Engel and Rodrigues (1989) considers the ex post rate of return on a currency which is defined as $y_{t+1} = (1 + i_t)S_{t+1}/S_t$, where S_t is the level of the nominal euro - dollar exchange rate, and is also be deflated by the relative price levels in each region. Then

$$E_t y_{t+1} = c + \rho_t B_t \lambda_t, \tag{3}$$

where $E_t(z_{t+1}z'_{t+1}) = B_t$ where z are fundamentals and risk factors considered to be associated with the euro currency and λ_t are factor loadings. The standard $I - CAPM$ imposes the restrictions that $\rho_t = \rho$ is constant.

In our formulation we impose a flexible auxiliary $I - CAPM$ formulation as the maintained hypothesis. It was previously seen that there are apparent discontinuities in the relationship with the bond markets and the euro. The usual linkages with equity markets most readily occurs via the discrete time intertemporal $CAPM$ which can be expressed as

$$E_t\left(R_{i,t+1} - R_{f,t+1}\right) = \beta Cov_t\left\{\left(R_{i,t+1} - R_{f,t+1}\right)\left(R_{M,t+1} - R_{f,t+1}\right)\right\} + \gamma Cov_t\left(R_{i,t+1}\Delta z_{t+1}\right)$$

where $R_{i,t+1}$ is the return on the ith asset between periods t and $t+1$, while $R_{f,t+1}$ denotes the risk free rate and is clearly known at time t, and Δz_{t+1} denotes the innovation in the state variables driving the system. The notation of Cov_t refers to the covariance conditioned on a sigma field of information available at time t. Cochrane (2005) shows that this is a valid discrete time approximation to the continuous time formulation of the general intertemporal $CAPM$ in continuous time of the Merton (1973) model. The fact that $\gamma \neq 0$ provides the point of departure form the standard $CAPM$ of Sharpe (1964) and justifies the use of multi factor terms which have tended to proliferate recent empirical domestic finance literature which focuses on describing the equity returns on domestic securities.

From using the law of iterated expectations the unconditional $CAPM$ can be interpreted as the above equation with essentially the two sources of risk that explain the average risk premiums. The static market risk premium from $CAPM$ is $\beta Cov_t\left(R_{i,t+1}, R_{M,t+1}\right)$ and simply implies that the price of an asset covaries positively with the market returns and earns a positive risk premium over the market rate. The estimate of the relative risk aversion parameter is generally considered that it should be in the range of 1 and 10.

The second term in equation (3) is intended to represent a source of risk that $\gamma Cov_t\left(R_{i,t+1}\Delta z_{t+1}\right)$ predicts future market returns. If the coefficient $\gamma > 0$ then an asset covaries positively with changes in the state variable and hence is positively correlated with future market expected returns.

In the context of our application based around the $I-CAPM$ the variable $(R_{i,t+1}-R_{f,t+1})$ is defined as the difference between the return on the Euro equity market and the Euro risk free rate. Similarly, the variable $(R_{M,t+1}-R_{f,t+1})$ is defined as the difference between the return on the US equity market and the US risk free rate. Possible further adjustments due to inflation differentials were not justified empirically. If the asset (exchange rate) does not provide a hedge against future negative shocks in the form of investment wealth then it offers low returns when future aggregate returns are also expected to be low. Hence a rational investor will only hold the asset (currency) if it offers an expected return in excess of the risk free rate.

When the model is pricing a set of assets in the cross section the state variable is positively correlated with the future aggregate returns and $\gamma > 0$.

One of the potential attractions of the international $CAPM$ is that it provides justification for the presence of state variables in terms of factors and as described by Fama (1991) is tantamount to providing a "fishing licence" for exploring sometimes arbitrary factors. However, from an international perspective when dealing with a currency this issue is perhaps easier to justify in terms of fundamentals and flows of capital between equity markets in different regions.

We consider estimation of the $I-CAPM$ from the European perspective with the $ESTOXX$ variable measuring returns on the domestic (European markets) and with the $S\&P500$ returns on the US equity market representing the "market" part of the $CAPM$ formulation. This seems a reasonable assumption given the size and importance of the US equity market within the world market. Also, this approach allows us to develop some simple theories to be developed between the US and Eurozone markets and hence relate to the euro-dollar exchange rate. In the following we also take the US T-Bill as the risk free rate to construct the $(R_{M,t+1}-R_{f,t+1})$ series. The equivalent or corresponding quantity for the Eurozone is less clear but we take the Euro bond rate since September 2004 for monthly data and since September 6, 2004 for daily data and German one month rate from January 1999 through August 2004 for monthly data and from January 1, 1999 through September 3, 2004 for the daily data analysis.

The other interesting aspect of the $I-CAPM$ formulation is that it allows for the presence of other factors, or state variables. Hence we use the VIX as an overall measure of risk and uncertainty and also the conditional variances of equity returns from both the US and European markets. This could potentially be extended to other factors. The momentum factor due to Carhart (1997) is popular in many domestic traditional Fama - French factor models. Our use of the TVP regression approach corresponding to variables including volatility can be interpreted as a form of momentum once it is included as a factor. To implement the concept of factor momentum we also report estimates of the $I - CAPM$ from daily data in the lower panel of figure 5 below the corresponding analysis of the monthly data for the similar model.

The application of the TVP methodology to the $I - CAPM$ generates parameter time paths that are represented in figure 5 for both monthly and daily data. The estimated β_t which measures how the price of euro denominated stocks covaries with the market returns and as expected is seen to earn a positive risk premium over the market rate for both the models from monthly and daily data. The estimated β_t is always positive and generally, but not always, less than one. The inclusion of the additional factors is generally not found to make any significant difference to the estimated β_t. There is the possibility of using other "factors" such as in the study by Coudert and Mignon (2013), who use credit default swaps as a proxy for sovereign debt risk. Overall, our results indicate a link between the US and European equity markets and also the euro-dollar exchange rate.

5 Conclusions

This paper provides a review of the econometric work that has accumulated so far on the Euro - US dollar exchange rate and to investigate possible reasons for the crises that have occurred. We concentrate primarily on the relevant asset markets that interact with the euro dollar exchange rate. Most of the empirical work in the paper uses the kernel weighted time varying parameter regression approach. This method has the advantage of allowing neat descriptions of variation in the structural parameters, some of which are directly associated with risk premium terms. We find evidence of substantial departures from parity conditions in the bond markets with significant positive and then negative excess returns over uncovered interest parity for some years after the 2008 financial crisis. We then estimate a time varying parameter model of a variant of the international $CAPM$, or $I - CAPM$. Much of the variation in the euro-dollar rate seems to be related to the variations in returns and their volatilities within both the US and European securities markets. Overall, our results indicate a link between the US and European equity markets and also the euro-dollar exchange rate.

One novelty of the work is that both daily and monthly econometric analysis is performed and the kernel weighted regression methods with time varying parameters turn out to be particularly useful in identifying periods of instabilities and crisis in the Eurozone. Some of the results suggest subsequent macro factor analysis to identify and investigate the causes of crises and instabilities.

There is the possibility of further work to identify changes in macro fundamentals or sources of exogenous risk which affect the $I - CAPM$ framework. The search for more factors is a possible area for future research.

6 Appendix: Kernel weighted regressions and the *TVP* Approach

There are many possible approaches at estimating β_t, the regression, and in the context of estimating the TVP regression in equation (7), the general TVP regression can be expressed as

$$y_t = x_t' \beta_t + u_t \tag{4}$$

with β_t being a bounded random walk, and in the context of equation and x_t' represents the explanatory variables. In general the kernel weighted regression estimator for β_t is

$$\hat{\beta}_t = \left(\sum_{j=1} w_{jt} x_j x_j' \right)^{-1} \left(\sum_{j=1} w_{jt} x_j y_j \right). \tag{5}$$

we use a similar approach by a kernel weighted regression and is a modification of the work on autoregressions by Giraitis, Kapetanios and Yates (2014) who assume that the TVP in an $AR(1)$ model follow a rescaled random walk, and where $\{a_t\}$ is a non stationary process which defines the random drift, and $-1 < \phi < 1$. However, in this context ϕ_t is a standardized version of a_t so that

$$\phi_t = \phi \frac{a_t}{\max_{0 \le k \le t} |a_k|} \tag{6}$$

It is assumed that $a_t = a_{t-1} + w_t$ which is a driftless random walk, where w_t is a stationary process with zero mean. If w_t is white noise then the process is identical to that of Cogley and Sargent (2005). Following Giraitis, Kapetanios and Yates (2014), it can be assumed that it is a general stationary process with possible slow hyperbolic decay in its autocorrelation function. They show that the coefficient process $\{\phi_t; t = 1, ..., T\}$ converges in distribution as T increases to the limit $\{\phi_t; 0 \leq \tau \leq 1\} \rightarrow_D \{\phi \tilde{W}_\tau; 0 \leq \tau \leq 1\}$. Giraitis, Kapetanios and Yates (2015) have extended this to the multivariate situation and following Baillie and Cho (2014), the approach for estimating the TVP is to take

$$\hat{\beta}_t = \frac{\sum_{t=1}^{H} K(\frac{t-k}{H}) y_t y_{t-1}}{\sum_{t=1}^{H} K(\frac{t-k}{H}) y_{t-1}^2} \tag{7}$$

where $K(\frac{t-k}{H})$ is a kernel and is a continuously bounded function. For example the Epanechnikov kernel with finite support and the Gaussian kernel with infinite support are potential candidates. Then,

$$H^{1/2}(1-\hat{\beta}_t^2)^{-1/2}(\hat{\beta}_t - \beta_t) \sim N(0,1) \tag{8}$$

Then

$$\hat{\beta}_t - \beta_t = \left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \sum_{j=1} w_{jt} x_j y_j - \beta_t = \left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \sum_{j=1} w_{jt} x_j \left(x_t' \beta_t + u_t\right) - \beta_t \tag{9}$$

If the bandwidth is $o_p\left(T^{1/2}\right)$, then the term $\left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \sum_{j=1} w_{jt} x_j x_j' \left(\beta_j - \beta_t\right)$ is negligible. Hence it is satisfactory to focus on the term $\left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \sum_{j=1} w_{jt} x_j u_t$. One expression for the estimator of the variance of the TVP if u_t is homoskedastic is given by

$$Var\left(\hat{\beta}_t\right) = \hat{\sigma}_u^2 \left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \sum_{j=1} w_{jt}^2 x_j x_j' \left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \tag{10}$$

where $\hat{\sigma}_u^2 = \frac{1}{T}\sum_{i=1}^{T}\left(y_t - x_t'\beta_t\right)^2$. If u_t is heteroskedastic then the covariance matrix of the TVP parameter estimates is given by

$$Var\left(\hat{\beta}_t\right) = \left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \left(\sum_{j=1} w_{jt}^2 x_j x_j' \hat{u}_t^2\right) \left(\sum_{j=1} w_{jt} x_j x_j'\right)^{-1} \tag{11}$$

Similarly, an extremum estimator of the form

$$\hat{\theta}_t = \arg\min_{\theta} \sum_{j=1}^{T} w_{jt} l_j\left(y_j; \theta\right) \tag{12}$$

will have a sandwich type estimator for the covariance matrix of the TVP parameter estimates given by

$$Var\left(\hat{\beta}_t\right) = \left(\sum_{j=1}^{T} w_{jt} \frac{\partial^2 l_j\left(y_j;\hat{\theta}_t\right)}{\partial\theta\partial\theta'}\right)^{-1} \left(\sum_{j=1}^{T} w_{jt}^2 \left(\frac{\partial l_j\left(y_j;\hat{\theta}_t\right)}{\partial\theta}\right)\left(\frac{\partial l_j\left(y_j;\hat{\theta}_t\right)}{\partial\theta}\right)'\right) \left(\sum_{j=1}^{T} w_{jt} \frac{\partial^2 l_j\left(y_j;\hat{\theta}_t\right)}{\partial\theta\partial\theta'}\right)^{-1}. \tag{13}$$

References

[1] Baillie, R.T., Bollerslev, T. and H.-O. Mikkelsen, 1996. Fractionally integrated generalized autoregressive conditional heteroskedasticity. Journal of Econometrics 74, 3–30.

[2] Baillie, R.T., Cho, D., 2014. Time variation in the standard forward premium regression: Some new models and tests. Journal of Empirical Finance 29, 52–63.

[3] Baillie, R.T., Kiliç, R., 2006. Do asymmetric and nonlinear adjustments explain the forward premium anomaly? Journal of International Money and Finance 25, 22–47.

[4] Beckmann, J., Belke, A. Kuhl, M. , 2011. The dollar-euro exchange rate and macroeconomic fundamentals: a time varying coefficient approach. Review of the World Economy, 147, 11–40.

[5] Bianco, M.D., Camacho, M., Quiros, G.P., 2012. Short-run forecasting of the euro-dollar exchange rate with some economic fundamentals. Journal of International Money and Finance 31, 377–396.

[6] Camarero, M., Ordonez, J. 2012. Nonlinear adjustment in the real dollar euro exchange rate: the role of the productivity differential as a fundamental. Economic Modelling 29, 444–449.

[7] Canarella, G., Miller, S.M., Pollard, S.K., 2014. Purchasing power parity between the UK and Germany: the euro era. Open Economies Review 25, 677–699.

[8] Carhart, M.M., 1997. On persistence in mutual fund performance. Journal of Finance 52, 57–85.

[9] Chen, H., Fausten, D.K., Wong, W., 2011. Evolution of the trans-atlantic exchange rate before and after the birth of the Euro and policy implications. Applied Economics 43, 1965–1977.

[10] Chinn, M.D., Moore, M.J., 2011. Order flow and the monetary model of exchange rates: Evidence from a novel data set. Journal of Money, Credit and Banking 43, 1599–1624.

[11] Cochrane, J. 2005. Asset pricing. Princeton University Press, Princeton, NJ.

[12] Coudert, V. , Mignon, V., 2013. The forward premium puzzle and the sovereign default risk. Journal of International Money and Finance, 32, 591–611.

[13] de Grauwe, P., 2016. The legacy of the eurozone crisis and how to overcome it. Journal of Empirical Finance. Forthcoming.

[14] Engel, C.M., Rodrigues, A.P., 1989. Tests of the international CAPM with time-varying co-variances. Journal of Applied Econometrics 4, 119–138.

[15] Evans, M., Lyons, R.K., 2002. Order flow and exchange rate dynamics. Journal of Political Economy 110, 170–180.

[16] Fama, E.F., 1991. Efficient capital markets: II. Journal of Finance 46, 1575–1617.

[17] Frankel, J.A., 1982. In search of the exchange risk premium: a six currency test assuming mean-variance optimization, Journal of International Money and Finance 1, 255–274.

[18] Frankel, J.A., Engel, C.M., 1984. Do asset demand functions optimize over the mean and variance of real returns? A six currency test. Journal of International Economics 17, 309–323.

[19] Giraitis, L., Kapetanios, G., Yates, T., 2014. Inference on stochastic time varying coefficient models. Journal of Econometrics 179, 46–65.

[20] Giraitis, L., Kapetanios, G., Yates, T., 2015. Inference on multivariate heteroskedastic time varying random coefficient models. QMUL Working Paper.

[21] Koedijk, K.G., Tims, B., van Dijk, M.A. 2004. Purchasing power parity and the euro area. Journal of International Money and Finance 23, 1081–1107.

[22] Lopez, C., Papell, D.H., 2007. Convergence to purchasing power parity at the commencement of the euro. Review of International Economics 15, 1–16.

[23] Lothian, J.R., Wu, L., 2011. Uncovered interest-rate parity over the past two centuries. Journal of International Money and Finance 30, 448–473.

[24] Merton, R. 1973. An intertemporal capital asset pricing model. Econometrica 41, 867-887.

[25] Sartore, D., Trevisan, L., Trova, M., Volo, F., 2002. US dollar/Euro exchange rate: a monthly econometric model for forecasting. European Journal of Finance 8, 480–501.

[26] Sharpe, W. 1964. Capital asset prices: a theory of market equilibrium under conditions of risk. Journal of Finance 19, 425–442.

TABLE 1. THE NUMBER OF TIME PERIODS THAT *UIP* IS REJECTED AT 95% LEVEL.

	Australia	Canada	Denmark	Japan	New Zealand
Number of months	0	25	200	0	0
that $\beta^{ub} < 1$	(0%)	(12.5%)	(100%)	(0%)	(0%)
Number of months	0	0	0	0	43
that $\beta^{lb} > 1$	(0%)	(0%)	(0%)	(0%)	(21.5%)
	Norway	Switzerland	UK	US	
Number of months	0	8	0	49	
that $\beta^{ub} < 1$	(0%)	(4%)	(0%)	(24.5%)	
Number of months	29	0	0	0	
that $\beta^{lb} > 1$	(14.5%)	(0%)	(0%)	(0%)	

Key: All results are based on the TVP kernel weighted regression model; where β^{ub} denotes the upper bound of the 95% confidence intervals, and β^{lb} denotes the lower bound of the 95% confidence intervals.

TABLE 2. ANALYSIS OF THE ESTIMATED BETAS FROM THE *TVP* KERNEL WEIGHTED REGRESSION

	In Figure 3		In Figure 4	
	Model 1	Model 2	Model 1	Model 2
Number of months	32	13		
that $\beta^{ub} < 0$	(16.1%)	(6.5%)		
Number of months	0	0		
that $\beta^{lb} > 0$	(0%)	(0%)		
Number of months	63	86	61	85
that $\gamma^{ub} < 0$	(31.7%)	(43.2%)	(30.7%)	(42.7%)
Number of months	0	0	0	0
that $\gamma^{lb} > 0$	(0%)	(0%)	(0%)	(0%)
Number of months	0	0	0	0
that $\delta^{ub} < 0$	(0%)	(0%)	(0%)	(0%)
Number of months	49	45	48	47
that $\delta^{lb} > 0$	(24.6%)	(22.6%)	(24.1%)	(23.6%)
Number of months		20		41
that $\theta^{ub} < 0$		(10.1%)		(20.6%)
Number of months		0		0
that $\theta^{lb} > 0$		(0%)		(0%)

Key: β^{ub}, γ^{ub}, δ^{ub}, and θ^{ub} denote the upper bounds of the 95% confidence intervals for the corresponding coefficients, and β^{lb}, γ^{lb}, δ^{lb}, and θ^{lb} denote the lower bounds of the 95% confidence intervals for the corresponding coefficients, in Figures 3 and 4.

TABLE 3. ESTIMATION RESULTS FROM *GARCH* AND *FIGARCH* MODELS FOR THE EURO-US DOLLAR DAILY SPOT RETURNS

$GARCH\,(1,1)$ model: $y_t = \mu + \varepsilon_t$, $\varepsilon_t \sigma_t^{-1} \sim i.i.d.N\,(0,1)$, $\sigma_t^2 = \omega + \alpha\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2$

$FIGARCH\,(1,d,0)$ model: $y_t = \mu + \varepsilon_t$, $\varepsilon_t \sigma_t^{-1} \sim i.i.d.N\,(0,1)$, $\sigma_t^2 = \omega + \beta\sigma_{t-1}^2$

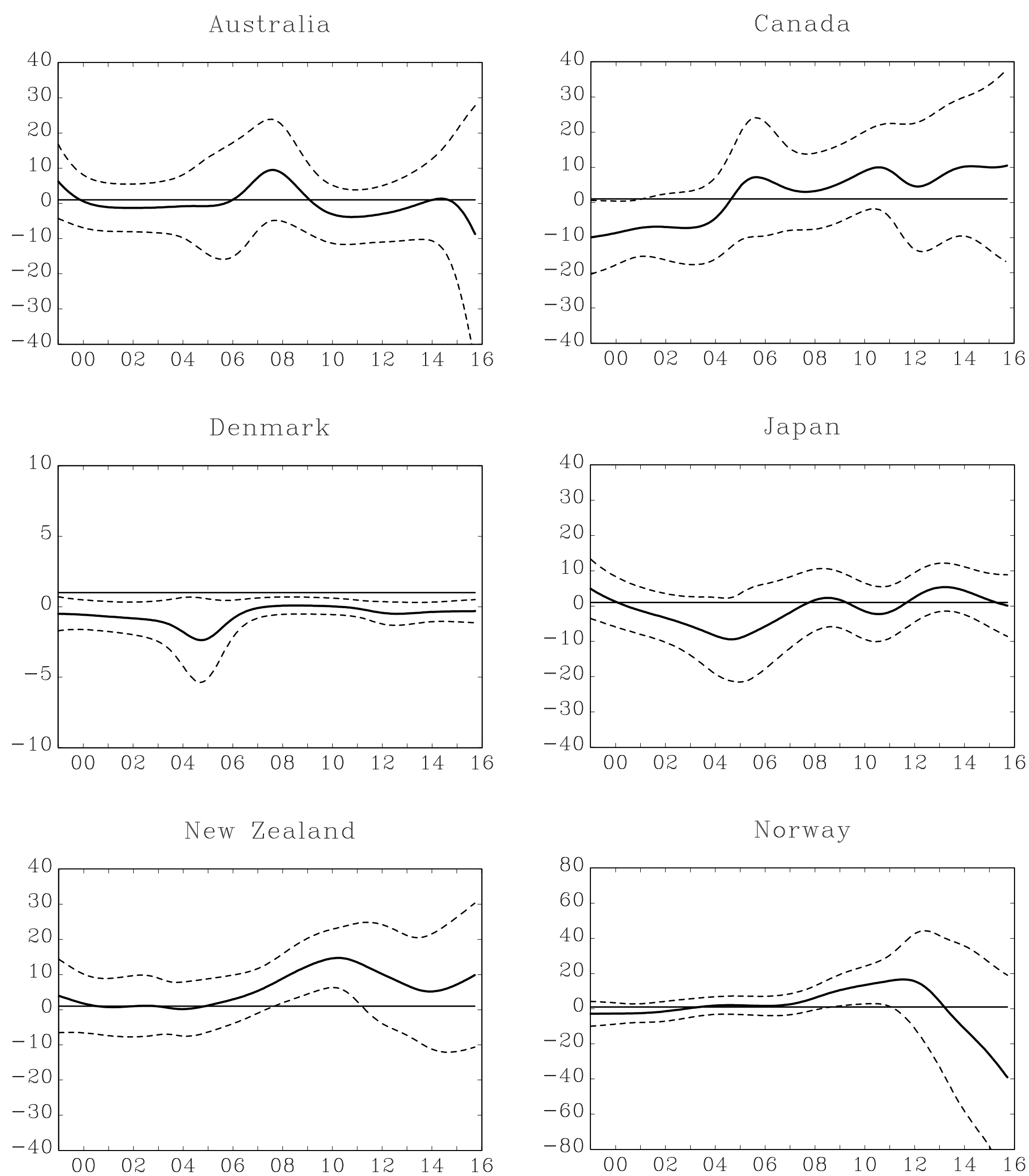
$$+ \left[1 - \beta L - (1-L)^d\right] \varepsilon_t^2$$

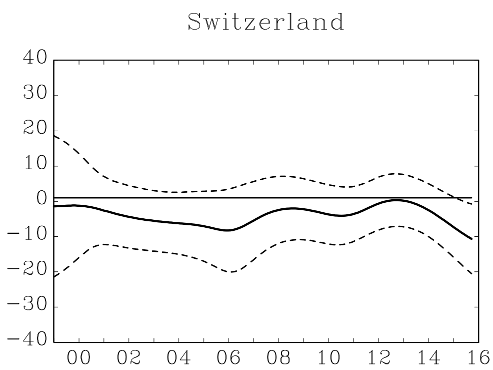
	Model	
Parameter	$GARCH(1,1)$	$FIGARCH(1,d,0)$
μ	0.045	0.031
	(0.085)	(0.171)
ω	0.104	4.691
	(0.066)	(1.002)
α	0.028	
	(0.003)	
β	0.970	0.285
	(0.004)	(0.031)
d		0.268
		(0.029)
Skewness	–0.034	0.024
Kurtosis	4.000	4.116
$Q(20)$	15.493	15.531
$Q^2(20)$	30.872	30.872

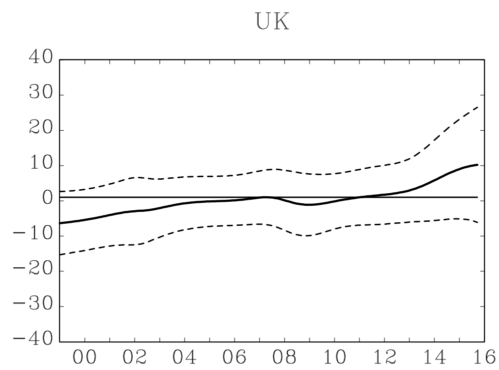
Key: The table reports the Quasi Maximum Likelihood Estimates ($QMLE$) for $GARCH(1,1)$ and $FIGARCH(1,d,0)$ models for the daily Euro-U.S. dollar spot returns from January 1, 1999 through September 30, 2015, for a total of 4,368 observations. The $QMLE$ are calculated under the assumption of conditional normality. Robust standard errors are reported in parentheses. The sample skewness and kurtosis for the standardized residuals are also reported. $Q(20)$ and $Q^2(20)$ refer to the Ljung-Box portmanteau tests for up to 20th-order serial correlation in the standardized and the squared standardized residuals, respectively.

FIGURE 1. SLOPE PARAMETER ESTIMATES FROM NEW *TVP* KERNEL WEIGHTED REGRESSIONS WHEN THE NUMÉRAIRE CURRENCY IS THE EURO. THE DASHED LINES REPRESENT 95% CONFIDENCE BANDS. THE STRAIGHT LINE REPRESENTS THE VALUE OF UNITY.

Model: $\Delta s_{t+k} = \alpha_t + \beta_t(i_t^* - i_t) + u_{t+k}$ where $k = 1$




Switzerland
40
30
20
10
0
−10
−20
−30
−40
00 02 04 06 08 10 12 14 16


UK
40
30
20
10
0
−10
−20
−30
−40
00 02 04 06 08 10 12 14 16

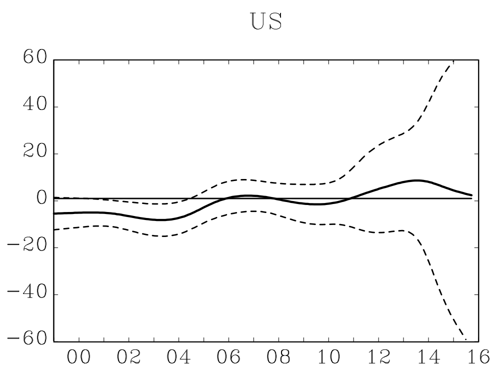

US
60
40
20
0
−20
−40
−60
00 02 04 06 08 10 12 14 16

FIGURE 2. SLOPE PARAMETER ESTIMATES FROM NEW *TVP* KERNEL WEIGHTED REGRESSIONS USING DAILY USD PER EUR. THE DASHED LINES REPRESENT 95% CIS. THE STRAIGHT LINE REPRESENTS THE VALUE OF UNITY.

Model: $\Delta s_{t+k} = \alpha_t + \beta_t(i_t^* - i_t) + u_{t+k}$ where $k = 21$

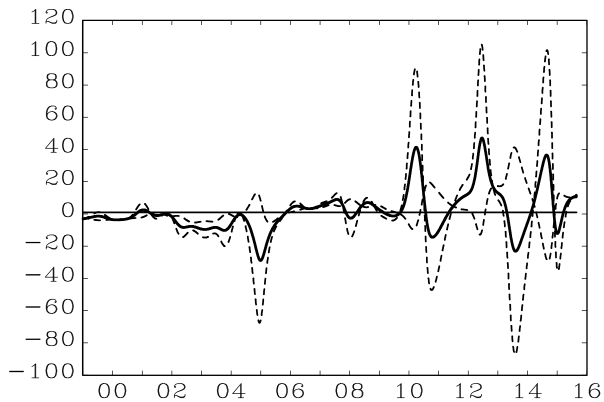
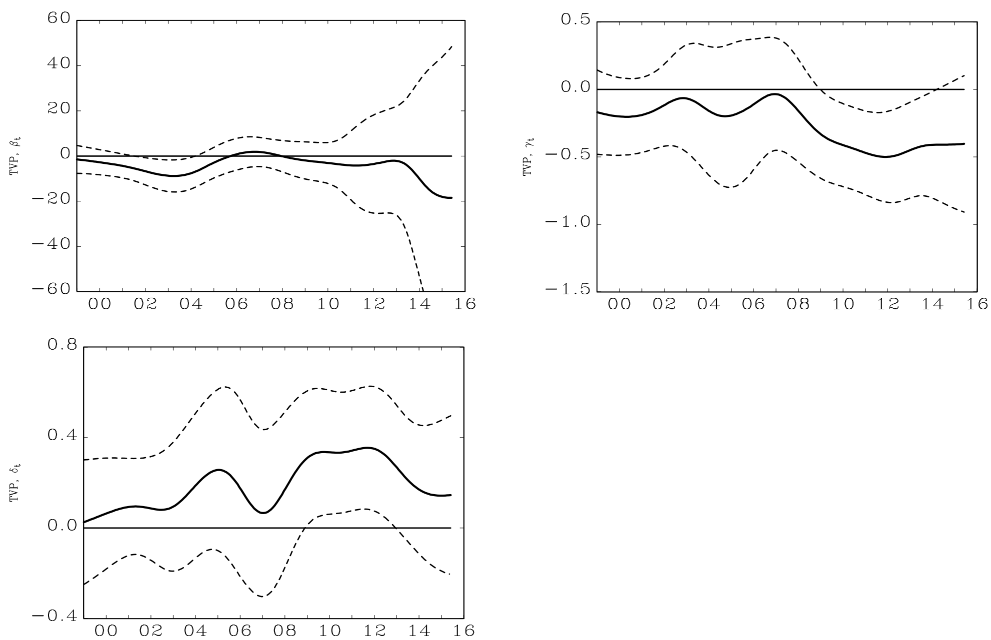


FIGURE 3. SLOPE PARAMETER ESTIMATES FROM NEW *TVP* KERNEL WEIGHTED REGRESSIONS WITH RISK FACTORS USING MONTHLY USD PER EUR. THE DASHED LINES REPRESENT 95% CIS.

Model 1: $\Delta s_{t+k} = \alpha_t + \beta_t(i_t^* - i_t) + \gamma_t S\&P_t + \delta_t ESTOXX_t + u_{t+k}$ where $k = 1$



Model 2: $\Delta s_{t+k} = \alpha_t + \beta_t(i_t^* - i_t) + \gamma_t S\&P_t + \delta_t ESTOXX_t + \theta_t \Delta VIX_t + u_{t+k}$ where $k = 1$

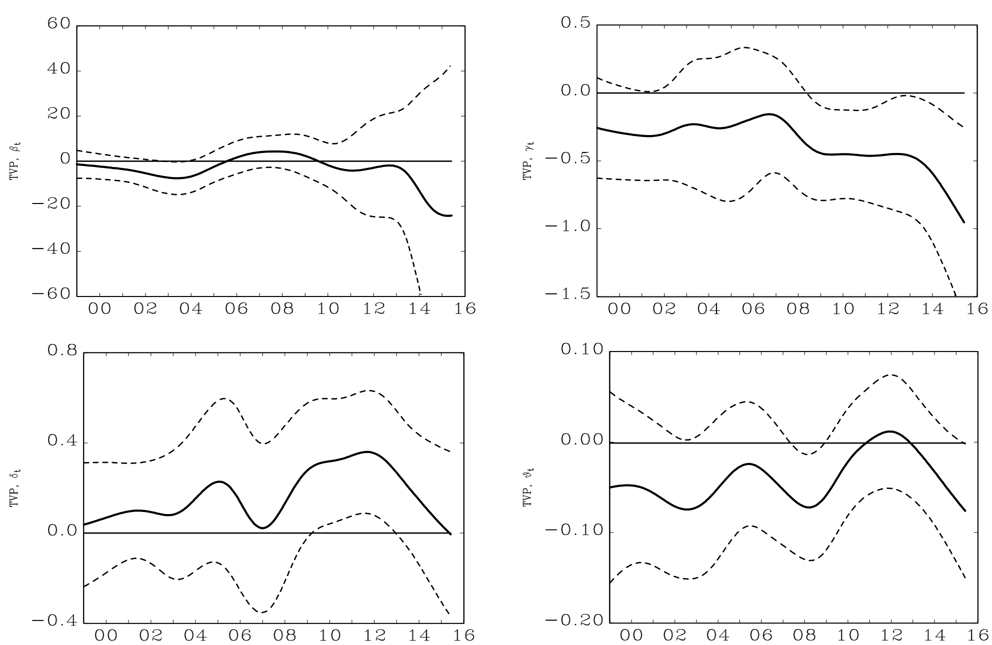
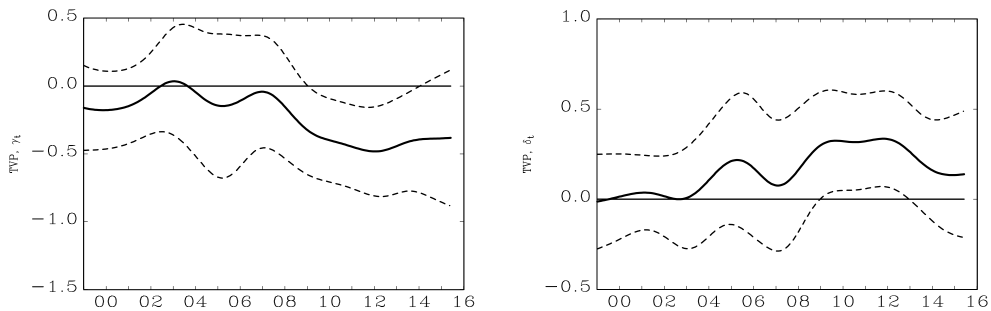


FIGURE 4. SLOPE PARAMETER ESTIMATES FROM NEW *TVP* KERNEL WEIGHTED REGRESSIONS WITH THE DEPENDENT VARIABLE BEING THE EXCESS SPOT RETURNS OVER UIP USING MONTHLY USD PER EUR. THE DASHED LINES REPRESENT 95% CIS. THE STRAIGHT LINE REPRESENTS THE VALUE OF ZERO.

Model 1: $[\Delta s_{t+k} - (i_t^* - i_t)] = \alpha_t + \gamma_t S\&P_t + \delta_t ESTOXX_t + u_{t+k}$ where $k = 1$



Model 2: $[\Delta s_{t+k} - (i_t^* - i_t)] = \alpha_t + \gamma_t S\&P_t + \delta_t ESTOXX_t + \theta_t \Delta VIX_t + u_{t+k}$ where $k = 1$

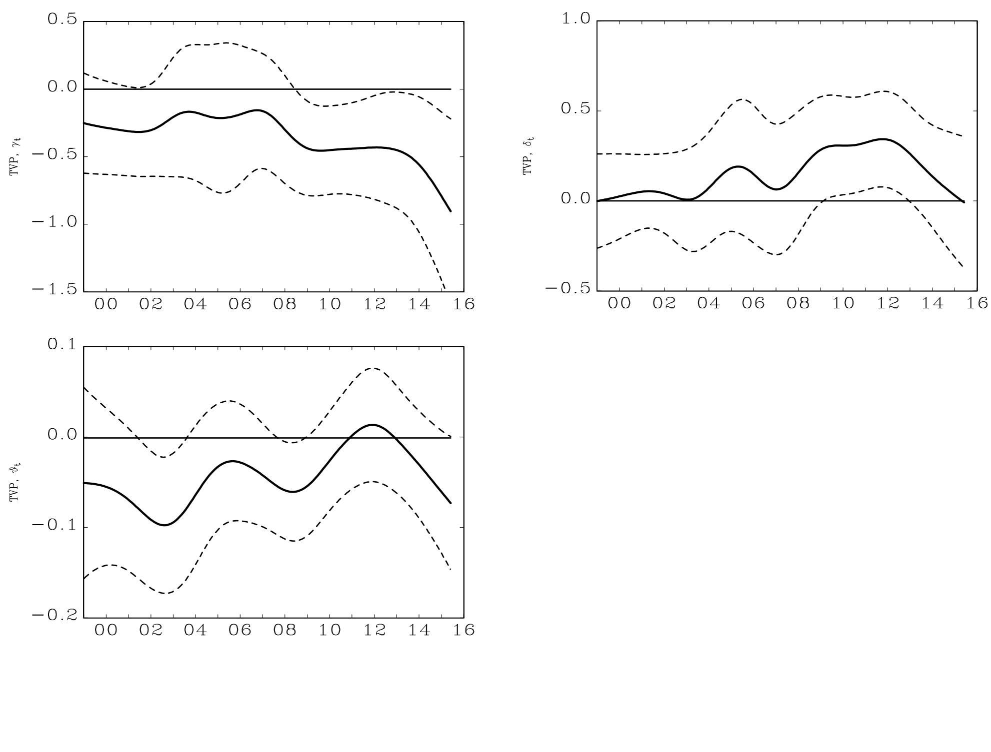
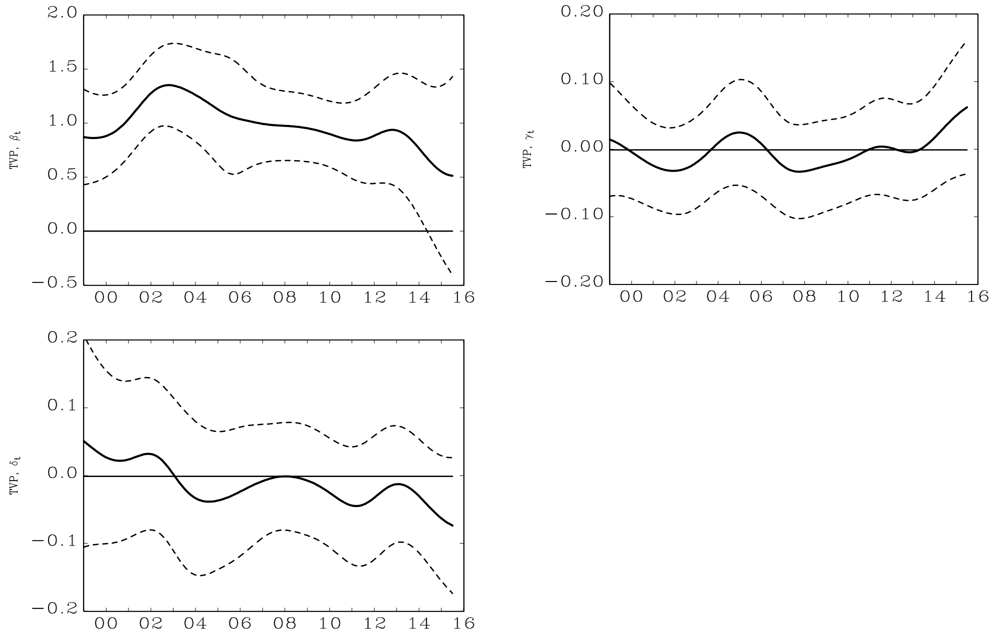


FIGURE 5. SLOPE PARAMETER ESTIMATES FROM NEW *TVP* KERNEL WEIGHTED REGRESSIONS USING MONTHLY AND DAILY USD PER EUR. THE DASHED LINES REPRESENT 95% CIS. THE STRAIGHT LINE REPRESENTS THE VALUE OF ZERO.

$$\text{Model: } E_t\left(R_{i,t+1}-R_{f,t+1}\right)=\beta_t Cov_t\left\{\left(R_{i,t+1}-R_{f,t+1}\right)\left(R_{M,t+1}-R_{f,t+1}\right)\right\}$$
$$+\gamma_t Cov_t\left(R_{i,t+1}\Delta\hat{\sigma}^2_{t+1}\right)+\delta_t Cov_t\left(R_{i,t+1}\Delta VIX_{t+1}\right)$$

where $\hat{\sigma}^2_{t+1}$ is the conditional variance of daily spot returns from the *FIGARCH* model.

(A) USING MONTHLY DATA



(B) USING DAILY DATA