New Evidence on the Forward Unbiasedness Hypothesis in the Foreign Exchange Market*

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Abstract

A large empirical literature has tested the unbiasedness hypothesis in the foreign exchange market using forward exchange rates. We amend the conventional testing framework to exploit the information in currency options, in an attempt to compare the test results obtained using forwards and options, and to assess the robustness of previous results reported by the literature. Applying our framework to a newly constructed data set for three major dollar exchange rates, we find that tests based on stationary regressions suggest that options provide biased predictions of the future spot exchange rate. Cointegration-based tests that allow for endogeneity problems arising from a potential omitted risk premium term are supportive of unbiasedness. We record strong similarities in the test results for forwards and options.

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Keywords: foreign exchange; market efficiency; unbiasedness; uncovered interest parity.

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1 Introduction

This paper revisits the unbiasedness hypothesis in the context of the foreign exchange (FX) market, one of the most researched and yet controversial hypotheses in the international finance literature. The unbiasedness hypothesis is related to the notion of FX market efficiency, as summarized by the uncovered interest rate parity (UIP) condition, which states that the expected exchange rate change should equal the current interest rate differential–or, in the absence of arbitrage, the forward premium (the difference between the forward and spot rates). Under UIP and in the absence of arbitrage (i.e. assuming that covered interest parity holds), the forward exchange rate provides an unbiased forecast of the future spot exchange rate, or, equivalently, the forward premium provides an unbiased forecast of the future change in the spot exchange rate–this is the key assertion of the unbiasedness hypothesis.[1]

The profession has long focused on investigating the relationship between changes in the exchange rate and the forward premium with less than satisfactory results. In a highly cited paper, Fama (1984) suggests that the expected change in the exchange rate is often inversely related to the forward premium, in stark contrast with UIP. This realization has spurred an enormous amount of research and produced a large spectrum of results which gave way to an extended list of possible explanations (e.g. Lewis, 1995; Flood and Rose, 1996; Engel, 1996; Berk and Knot, 2001; Chinn and Meredith, 2004). In general, however, tests of the unbiasedness hypothesis for different currency pairs and time periods gave further credit to Fama's results, which are now considered a stylized fact (Froot and Thaler, 1990), giving rise to the "forward bias puzzle," one of the central puzzles in international finance.

In this paper we re-examine the unbiasedness hypothesis, by changing vehicles of forming predictions about the future spot exchange rate. Specifically, we switch from the forward to the options market. Using data from the Philadelphia Exchange (PHLX), we construct a synthetic forward contract, made of currency options, which we call "option equivalent contract" and substitute it for the standard forward contract in the analysis of unbiasedness. We compare our results with the results obtained using forward contracts. Throughout this study, we combine conventional methods (the typical UIP regression first used by Fama) with the latest advances in the relevant literature on cointegration-based tests for unbiasedness, so as to present a thorough analysis for both forward- and option-based unbiasedness tests.

This approach yields several original additions to the relevant literature. First, we provide evidence on

[1] In our terminology in this section, tests of UIP are essentially interchangeable with tests of the unbiasedness hypothesis–that is the coefficient on lagged interest differentials or forward premia (lagged forward rate) in regressions of current exchange rate changes (current exchange rate level) is unity. This is somewhat loose, in that UIP is a sufficient but not necessary condition for unbiasedness, as discussed later in the paper.

the empirical validity of the unbiasedness hypothesis using currency options in a novel testing framework, complementing the conventional testing procedure which was, until now, restricted to forward markets. Second, our unique data set allows us to create a bridge between the over-the-counter (OTC) market for forward contracts and the organized exchange for options contracts. Our research directly compares the two derivatives markets, the forward and the options market, in terms of the statistical properties of the resulting contracts. Third, our empirical work allows us to assess whether the bias puzzle recorded in the literature to date is "forward specific" or a problem of a more general nature which is likely to be pervasive in other derivatives contracts.[2]

Our results provide several useful insights. The methodology used to produce the option equivalent contracts results into a synthetic forward which may be compared to the conventional forward contract, from which it differs in terms of contract specifications (e.g. maturity, expiry, trading specifications). The resulting option equivalent and the forward rate exhibit striking similarities in terms of both statistical properties and test results relating to the unbiasedness hypothesis. Overall, when using conventional tests based on stationary regressions we conclude that there appears to be an "option bias puzzle," reinforcing the case for the well documented "forward bias puzzle." However, more powerful cointegration tests designed to allow for the presence of a potential risk premium–which does not, *per se*, preclude unbiasedness–show ample support for unbiasedness for both forwards and options. The results are found to be robust to a variety of different departures from the core analysis, including the frequency of the data and the maturity of the derivatives contract.

The structure of the paper is as follows. In Section 2 we present a brief review of the literature on testing the unbiasedness hypothesis and discuss our motivation for the use of options in this context. In Section 3 we describe our data set and provide details related to the construction of the synthetic forward for our purposes. Section 4 presents the empirical results of our core analysis, while Section 5 reports robustness checks of the core results. Section 6 briefly summarizes and concludes. A number of robustness results are tabulated in Appendices A and B.

2 Testing the unbiasedness hypothesis using options

In this section we briefly review the enormous literature testing the validity of UIP and the unbiasedness hypothesis in the FX market, which has led to mixed results. Specifically, on the one hand tests based on

[2] In other words, we address the question whether the puzzle is likely to be caused by some specific characteristics of the forward market (e.g. the specific way that agents in this market form predictions about the future spot rate), or it is pervasive in other derivatives markets as well.

stationary regressions (e.g. research following Fama, 1984) have recorded that the forward premium is not an unbiased predictor of the future rate of depreciation, and in fact there is a forward bias such that the forward premium is generally inversely related to future movements in the exchange rate. On the other hand, more recent cointegration-based tests that allow for endogeneity problems caused by a potential unobserved risk premium provide some supportive evidence for the forward rate unbiasedness hypothesis (Barnhart, McNown and Wallace, 1999; Maynard, 2003), although cointegration studies provide, overall, mixed results.[3] We then describe how we amend the conventional unbiasedness tests by substituting the forward exchange rate with a suitably constructed proxy for the market expectation, based on information embedded in options contracts. We term such proxy the "option equivalent."

2.1 Conventional tests of forward rate unbiasedness

UIP purports that the FX gain from holding one currency instead of another–the expected exchange rate change–must be offset by the opportunity cost of holding funds in one currency rather than the other–the interest rate differential:

$$\Delta_k s^e_{t+k} = i_{t,k} - i^*_{t,k} \tag{1}$$

where s_t denotes the logarithm of the spot exchange rate (domestic price of foreign currency) at time t; i_t and i^*_t are the nominal interest rates available on similar domestic and foreign securities respectively (with k periods to maturity); $\Delta_k s^e_{t+k} \equiv s^e_{t+k} - s_t$; and the superscript e denotes the market expectation based on information at time t. In its simplest form, FX market efficiency can be reduced to a joint hypothesis that FX market participants are, in an aggregate sense, (a) endowed with rational expectations and (b) risk-neutral. The hypothesis can be modified to adjust for risk, so that it then becomes a joint hypothesis of a model of equilibrium returns (which may admit risk premia) and rational expectations.

In practice researchers investigate UIP with the aid of Covered Interest Parity (CIP), the most common no-arbitrage relationship in the context of foreign exchange. CIP finds its mathematical representation in the form: $f^k_t - s_t = i_{t,k} - i^*_{t,k}$, where f^k_t is the logarithm of the k-period forward rate (i.e. the rate agreed now for an exchange of currencies k periods ahead). Should CIP not hold at a point in time, profitable opportunities would emerge, which would induce trade in opposite directions resulting to their

[3] The lack of consensus in empirical research on forward rate unbiasedness is well characterized by Engel (1996, p. 141) as follows: 'To summarize [...] some have found [the future exchange rate and the current forward rate] are cointegrated with cointegrating vector $(1, -1)$; some have found they are cointegrated but not with a cointegrating vector $(1, -1)$; and some have found that they are not cointegrated. These conflicting results hold on tests for the same set of currencies.'

elimination.[4]

Assuming that CIP holds, UIP can be re-written as $\Delta_k s^e_{t+k} = f^k_t - s_t$, i.e. the forward premium (or forward discount) $f^k_t - s_t$ should equal the market expectation of the exchange rate change $\Delta_k s^e_{t+k}$; and $s^e_{t+k} = f^k_t$, i.e. the forward rate should be an unbiased predictor of the future spot rate. A test of this hypothesis involves regressing the exchange rate change on the lagged forward premium and, following much previous literature, we shall refer to this regression as the 'Fama regression' (Fama, 1984):

$$\Delta_k s_{t+k} = \alpha + \beta \left(f^k_t - s_t \right) + \varepsilon_{t+k} \tag{2}$$

where, under UIP, $\alpha = 0$, $\beta = 1$, and ε_{t+k} is a white noise error. The empirical results from estimating regression (2) have led to strong rejections of UIP and, hence, FX market efficiency (e.g. see the references in the survey of Hodrick, 1987; Lewis, 1995; Engel, 1996). While α is generally close to zero and often statistically insignificant, β is estimated to be far from its theoretical value of unity and it is often found to be negative and statistically significantly different from zero. Indeed, it is a stylized fact that estimates of the slope parameter β are generally closer to minus unity rather than plus unity (Froot and Thaler, 1990). The negative value of β is the central feature of the forward bias puzzle, one of the most robust puzzles in international finance, which remains unexplained even with 20 years of hindsight since the work of Fama (1984).[5]

Another strand of the literature, building on ideas initially put forth by Fama (1984), and further elaborated by Liu and Maddala (1992) and Barnhart, McNown and Wallace (1999), claims that the conventional Fama regression is invalidated, due to problems of endogeneity, which may result from the appearance of an unobserved risk premium. Specifically, note that the vast majority of studies in this context estimate the Fama regression using ordinary least squares (OLS). This can be problematic in the presence of an omitted risk premium in the Fama regression, in which case OLS would yield biased and inconsistent estimates of β due to a simultaneity problem (Fama, 1984; Liu and Maddala, 1992; McCallum, 1994). Recently, Barnhart, McNown and Wallace (1999) have formally shown that two

[4] Extensive empirical evidence provides support to the validity of CIP (Frenkel and Levich, 1975, 1977; Levich, 1985; Frankel and MacArthur, 1988; Taylor, 1987, 1989; for a survey of this evidence, see e.g. Sarno and Taylor, 2003, Ch. 2). Note that, unlike CIP, UIP is not an arbitrage condition since one of the terms in the UIP equation, namely the exchange rate at time $t+k$, is unknown at time t and, therefore, non-zero deviations from UIP do not necessarily imply the existence of arbitrage profits due to the FX risk associated with future exchange rate movements.

[5] Exceptions include Bansal and Dahlquist (2000), who document that the forward bias is largely confined to developed economies and to countries for which the US interest rate exceeds foreign interest rates; Bekaert and Hodrick (2001), who, paying particular attention to small-sample distortions of tests applied to UIP and expectations hypotheses tests, provide a 'partial rehabilitation' of UIP; and Flood and Rose (2002), who report that the failure of UIP is less severe during the 1990s and for countries which have faced currency crises over the sample period investigated.

conditions are needed for the simultaneity problem to arise: (i) the forward rate must be a function of an unobservable omitted variable, such as predictable excess returns; (ii) the term containing the forward rate in the estimated regression must be stationary or, if nonstationary, can be normalized to a stationary variable. Under these conditions, Barnhart, McNown and Wallace document the severity of this problem in a variety of spot-forward regressions, concluding that most common tests of unbiasedness are non-informative in the presence of simultaneity. Failure to properly account for these factors results into correlation between the forward premium and the error term, which induces the bias to assume values bigger than unity, therefore driving the OLS estimate of β towards negative values. This simultaneity problem renders the estimates from the Fama regression, and other derivative formulations of UIP tests, biased and inconsistent.

The proposed remedial methodology is to carry out a cointegration analysis involving the level of the spot exchange rate and the lagged forward rate (an unbiased predictor of spot rate under the unbiasedness hypothesis). Specifically, Barnhart, McNown and Wallace (1999) formally demonstrate that a forward unbiasedness test that is immune from the endogeneity problem involves two steps: first, testing for the existence of a cointegrating relationship of form $[1, -1]$ between s_{t+k} and f_t^k; second, if this cointegrating relationship holds, then a test for forward unbiasedness involves testing for residual correlation (both own and cross-currency correlation). Forward unbiasedness holds if this exact cointegrating relationship is validated by the data *and* the stationary residuals are white noise, suggesting that no incremental information can be added using available information at time t. In contrast to the results from estimating the Fama regression (2), the evidence from forward unbiasedness tests based on a cointegrating framework, which effectively allows for a potential risk premium term in the relationship between forward rates and future spot rates, and subsequent residual tests lends some support to the hypothesis that the forward rate is an unbiased predictor of the future spot exchange rate (e.g. Liu and Maddala, 1992; Barnhart, McNown and Wallace, 1999).[6]

In essence, the literature provides mixed evidence on the validity of the forward unbiasedness hypothesis. Studies employing the Fama regression provide robust evidence of a forward bias puzzle (β different from unity and often negative or statistically insignificant), whereas some recent studies based on a cointegrating framework to allow for the endogeneity problem discussed above suggest that the forward rate is unbiased predictor of the future spot rate. All of this evidence is based, however, on one specific derivatives contract, namely the forward exchange rate, in order to proxy the market expectation

[6] For interesting related results in the context of cointegration in other derivatives markets, see Kellard, Newbold, Rayner and Ennew (1999).

of the future spot rate.

2.2 Using options to test the unbiasedness hypothesis

In this paper we endeavour to find a different, and yet simple, path to test the unbiasedness hypothesis. Our swift of focus on FX options, as predictors of the future spot exchange rate, not only bears some plausible intuition, but also acts as a robustness check for the previous results documented in the literature based on forward contracts.

Our reasoning for choosing the options market resides in our effort to extract information from a different FX derivatives market than the conventional forward market, yet bearing an extensive involvement in the FX market practices. Apart from trading and contract setting conventions, which induce differences between the forward and options markets (briefly presented in Section 3), we offer two intuitions as to why options may contain somewhat different information from forwards. In analyzing options, one should bear in mind that, intuitively, an option contract represents a bet that the price of the currency examined will be above or below a certain level. Investors who believe that the price will rise buy call options and those who believe that the price will fall buy put options. We are interested in investigating whether the different tenets of the two markets, as described below, would induce or compel different betting behavior by agents.

First, it may seem that options contain different information than forward contracts due to the flexibility (of whether and/or when) to exercise an option.[7] In the case of options, the strike price is a mere reference point as the investor targets a wider range of values above or below the actual strike price. On the contrary, in forward contracts, where no such flexibility exists, the settlement price is the exact betting price and the investor aims for a final result as close to that price as possible.

Second, option contracts at a specific time t can have various strike prices: as the degree of moneyness of the contract changes, contracts with new strike prices are introduced to always ensure the presence of put and call contracts. Further to that, tailor-made contracts are also introduced for different strike prices according to the needs of the clients. Therefore, it is possible to have a whole distribution of strike prices for otherwise identical contracts at each time t, which can potentially better capture the expectations of the market. This again contrasts with forward contracts, for which researchers are only

[7] American options contracts are perhaps the clearest example, because they include both the downward insurance to the investor that European contracts have (in case of an unfavorable outcome the investor leaves the option unexercised and only loses the premium paid to acquire the option) plus an additional time value parameter, which comes from the flexibility to exercise on or before expiry (the investor has the advantage of being able to wait for the most appropriate time to exercise the option).

presented with a single forward rate at time t from available databases. This property adds a further dimension to option contracts, namely the distribution of strike prices, which adds to the widely used term structure of contract prices.

The above traits present us the opportunity to test the unbiasedness hypothesis by applying a different instrument. We create a synthetic forward contract (termed the options equivalent, or simply o). Our aim is to re-examine forward unbiasedness by: (a) using all the relevant conventional methods based on both the Fama regression and on cointegration analysis; (b) presenting a thorough investigation of both forward and options by undertaking several robustness tests to investigate whether the two derivatives markets yield similar results in terms of portraying investors' expectations; (c) determining whether the forward bias puzzle is indeed specific to the forward market or a more general feature of FX markets.

2.2.1 Creating the synthetic forward

In order to compare the forward and options markets, we begin from calculating an option measure that is equivalent to the forward rate, i.e. a synthetic forward contract. To this end, we combine the arbitrage conditions of both the forward and the options market. The most prominent arbitrage condition in the options market is the Put-Call Parity (PCP) condition, which establishes a relationship between European put and call option prices (e.g. Stoll, 1969; Merton, 1973; for the case of options in the FX market see Grabbe, 1983). More specifically, using capital letters to relate to levels (i.e. no longer logs of values), PCP suggests that buying a call ($-C$) and selling a put ($+P$) with the same strike price (K) and for the same underlying asset (in our case an exchange rate (S)) must yield exactly the same payoff to an investor as a synthetic long forward; i.e., given a domestic interest rate i and a foreign interest rate i^*, the strategy involves borrowing $K/(1+i)$ domestic, purchasing foreign currency and investing $S/(1+i^*)$ worth of foreign currency abroad (Levich, 2001). Therefore:

$$-C + P = \frac{K}{1+i} - \frac{S}{1+i^*} \tag{3}$$

or, equivalently

$$C - P = \frac{S}{1+i^*} - \frac{K}{1+i} \tag{4}$$

for the reverse strategy.

Although only a thin branch of the literature has tested the validity of this no-arbitrage condition, the available empirical evidence suggests that observed violations occur rarely and do not last long, supporting the assumption of no arbitrage (Shastri and Tandon, 1985, 1986; Bodurtha and Courtadon, 1986, 1987; El-Mekkaoui and Flood, 1998). We combine CIP and PCP to get the so called Put-Call Forward (PCF) parity relation (Grabbe, 1983):

$$\frac{F-K}{1+i^*} = (C-P). \tag{5}$$

A simple reparametrization of this parity relationship gives us the desired synthetic forward contract price, termed the option equivalent O, which is essentially a synthetic forward contract made of options:

$$O \equiv F = K + (C-P)(1+i^*). \tag{6}$$

This representation relates the forward price to the price of put and call option contracts on the same strike price. Note that this specific formulation applies to European options, while the relevant equation for the option equivalent would hold with inequality for American options.

Intuitively, equation (6) suggests that the option equivalent to the forward rate is the amount by which the final price will either exceed (if the call contract is exercised) or fall below (if the put contract is exercised) the strike price. Further details on the construction of the option equivalent are provided in the following section. Defining o as the log-option equivalent, the Fama regression may be written in terms of a link between the spot rate change and the option premium as follows:

$$\Delta_k s_{t+k} = \alpha + \beta \left(o_t^k - s_t\right) + \eta_{t+k} \tag{7}$$

which has the usual interpretation, i.e. $\alpha = 0$, $\beta = 1$ and η_{t+k} is a white noise error under FX market efficiency. Note that in equation (7) we use α and β to denote the constant term and the slope parameter respectively to ease the comparison with the corresponding parameters in the conventional Fama regression (2).[8]

3 Data

3.1 Sources

The data set we employ in the empirical work consists of weekly spot, forward, synthetic forward and interest rate (eurocurrency) data. We employ data at weekly frequency, which had to be carefully constructed from intraday data. The sample period spans from 3 January 1986 to 31 December 2003.

[8] Clearly, in the case of American options, one would be dealing with an inequality in equation (3) for PCP and the resulting option equivalent in equation (6) would, therefore, also hold with inequality.

The synthetic forward was constructed from intraday data on options, provided by the Philadelphia Exchange (PHLX), the main currency options exchange in the US according to the 2004 Bank for International Settlements (BIS) *Survey on Foreign Exchange and Derivatives Market Activity*.[9] The PHLX has kindly made available to us the full trading history tape, which consists of all recorded transactions on standardized currency options contracts from 1986 to 2003. Specifically, the tape records the characteristics of all put and call options being traded (underlying currency, option premium, strike price, expiry date and number of contracts trading at the specific price) as well as the spot price of the underlying currency with time precision to the nearest second. Bid and ask spreads for both spot and option prices are being recorded non-continuously and are, therefore, not being used for reasons of consistency within the same series and with the forward data. Thus, we use the mid-point for both spot and option prices, which is consistent with the literature testing the forward unbiasedness hypothesis. We use data from 02:30am to 02:30pm (Philadelphia time), which includes the hours of main trading activity throughout the sample.[10] We focus on the most actively traded contracts, namely American contracts with mid-month expiry, for three major dollar exchange rates against the UK sterling, the Japanese yen and the Swiss franc (GBP, JPY and CHF).

Although the PHLX also trades European contracts, we investigate American contracts because they are by far the most heavily traded. We include all available trades, without excluding cases of potential early exercise. Some authors adopt the practice of excluding such cases in the context of testing the validity of PCP (Shastri and Tandon, 1985, 1986; Bodurtha and Courtadon, 1986, 1987). However, we prefer using all of the observations because our focus is not on testing PCP and because excluding trades that are deep in the money would distort the intraday distribution from which we construct our time series since it would skew it to the left.

It is also worth noting that, although researchers have studied subsets of the PHLX data in previous papers for other research purposes, the current paper analyzes the longest PHLX span ever considered in empirical work, with the full tape consisting of about 1,800,000 intraday observations for each time series examined.

[9] The difference between the OTC market and the organized exchanges is significant. OTC markets are decentralized markets that provide flexibility in option contracts (tailor-made contracts), customizing their specifications. In contrast, organized exchanges are largely centralized markets, offering standardized contract specifications and market conventions. Organized exchanges represent a smaller venue for currency option trading, compared to the OTC market, although the former is growing at a somewhat stronger pace (BIS, 2004).

[10] During the history of currency options, which only begun in 1982, the PHLX has experimented with various timing schedules for its operations, in response to demand from different world sectors, resulting to an around-the-clock trading session in 1990. However, lately its operations have been scaled back and its current currency option trading hours are from 02:30am to 02:30pm (Philadelphia time).

Data on 1- and 3-month forward contracts, spot rates and interest rates (eurocurrency rates), at the daily frequency, for the same set of currencies and sample periods as above, were provided by the BIS. These data were converted to weekly frequency, in such a way as to match the dates available for our weekly forward equivalent series.

3.2 Data details and manipulation

Our aim is to transform the intraday data on options contracts into a weekly series of synthetic forward contracts. We focus on a specific day of the week, namely Friday (the day of the contracts' expiry), thus creating weekly time series where each weekly observation corresponds to the last trading day of the week.[11] The intraday option equivalent was constructed using equation (6) for the intraday data on each Friday, matching put and call contracts with identical contract specifications for trades occurring within 5 minutes from each other. Then, in order to move from intraday (intra-Friday) data to weekly time series, and given that such an option equivalent is constructed for the first time, we adopted various approximations for the representative weekly quote. First, we employed the last trade of each Friday (*Last Trade*), to conform to the forward practices in the literature. Second, we constructed the Friday's average, that is the mean of the distribution of the intraday synthetic forward on each Friday (*Average*). A third measure was constructed from the mean of at-the-money contracts (*ATM*), thus screening what are typically the most frequently traded contracts. Fourth, we considered the median of the distribution of the intraday synthetic forward (*Median*). Lastly, we employed the volume of trade for each contract as a weight and calculated the weighted Friday's average of synthetic forward contracts (*W. Average*). Note that we have applied the above techniques to construct both the synthetic forward and its respective spot rate, in order to match the option equivalent as closely as possible with the corresponding spot exchange rate on each Friday.

Given the novelty of the resulting data set, this seemingly simple process was confronted with several challenges. A drawback in our data set was the decline in trading after 1995, which became apparent around 1997 and onwards. Although a straightforward explanation for such decline in trading is not provided by the PHLX, possible explanations gathered through our telephone interviews with PHLX managers include the gradual swift of focus to electronic trading and, most predominantly, the investors' increasing preference for the United Currency Options Market (UCOM), a November 1994 innovation of the PHLX, which offers the possibility to customize the contract specifications.[12] The above drop in

[11] In cases of no data on Fridays (e.g. due to public holidays) we use the immediately preceeding day within the same week.

[12] Up to that point, the PHLX only offered standardized contract specifications and market conventions, i.e. contracts

observations weighted on our effort to construct weekly estimates from a comparable number of intraday observations. For robustness, we therefore also created a different set of series where we removed outliers, defined as observations of the option equivalent in the 5th and 95th quantiles of the distribution. We use the outlier-removed sample to check the robustness of our empirical results; as discussed in our empirical work below, our results are qualitatively identical for these two different sets of data.

Another challenge we faced involved matching the conventions of the standardized (options exchange) market with the customized (OTC) forward market in terms of maturities and expiry dates. In the forward market, the expiry date of the contract can be on any day of the month, for contracts of any conventional maturities used in the forward market. The time to maturity becomes an immediately observable feature of forward contracts. On the contrary, the PHLX offers only two expiry days per month–namely mid-month and month-end expiries respectively. Therefore, the observable features of the trades on the PHLX are the expiry date and, consequently, the type of contract–recall that we only use mid-month contracts in this paper. This implies that for the forward market we can observe the maturity of the contract expiring on each day, whereas in the PHLX system options contracts of all maturities expire on a specific day of each month.[13] Naturally, these differences reduce the comparability of forward and option equivalent rates.

For the construction of the 1- and 3-month synthetic forward contract we fixed the expiry date and gathered all the relevant contracts that have been traded 1 month (4 weeks) and 3 months (12 weeks) ahead. Every time the expiry date was reached, the contracts of the expiry date entered the new cycle with a next expiry date. Therefore, at a specific time all contracts selected will expire on the same specific date, although the contracts might have begun trading at different points in the past. As a result, our 1- and 3-month synthetic forward contract is different from our 1- and 3-month conventional forward contract in that the former has a specific expiry date, based on the expiry cycle of the options, as specified by the PHLX, and includes all contract maturities trading within these dates, whereas the latter has fixed time to maturity but can expire any day of the month. This feature of the synthetic forward allows us the flexibility of assigning different values for k in the Fama regression given by equation (2) when we use the option equivalent. Namely, k can take the values 4, 8, 12, 16, 24, 36, 52 (weeks), in

that specify the currency pair traded, the contract size, strike price intervals, expiration dates, price quoting and premium settlement. UCOM increased flexibility by introducing customized currency options. This offered a choice to investors over all aspects of a currency option trade (exercise price, selection of currency pairs, premium quotation as either units of currency or percent of underlying value, and customized expiration dates of up to two years).

[13] The PHLX has standardized expiry dates, by setting specific expiry conventions. Mid-month contracts expire only on the first Friday following the third Wednesday of the expiry month, and month-end contracts expire on the last Friday of the month. The contracts trade on a fixed-months quarterly cycle (March, June, September, December and the two months following the current month).

contrast with the forward rate which in the literature is typically used for 4, 8 or 12 (weeks). In this paper, however, we study the 4-, 8- and 12-week contracts for our synthetic forward contracts to make a tighter comparison with the relevant literature.

The resulting series of interest are as follows. For tests based on forward contracts, the data set includes the logarithm of the spot exchange rate, s_t and the logarithm of the 1- and 3-month forward exchange rates, f_t^4 and f_t^{12} respectively, at weekly frequency–specifically, end-of-the-week prices. For tests based on options contracts, the series of interest consist of the logarithm of the five different definitions for the synthetic forward at time t, $\left(o_{t,j_h}^k\right)$ and the respective spot $(s_{t,h})$, where the subscript $j =$ *ATM, Average, Last Trade, Median* and *W. Average* corresponds to the five different methods of data construction described above; the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively; and $k = 4, 8, 12$ is the maturity of the contract in weeks. This data set provides a variety of distinct sets of spot and synthetic forward rates for each of the three exchange rates we examine. Given the vast amount of results we obtained, the core of the empirical work is based on weekly data for s_t and f_t^4 for forward-based tests, and $s_{t,h}$ and o_{t,j_h}^4 for options-based tests, while we shall use the remaining data in our robustness analysis.

4 Empirical results

4.1 Preliminary data analysis[14]

As a preliminary exercise we compared the sample moments of weekly spot and forward rate changes, changes in the synthetic forward rates, as well as the forward premium, $\left(f_t^k - s_t\right)$ and the option premium, $\left(o_{t,h}^k - s_{t,h}^k\right)$. For the individual series in first difference (spot, forward and option equivalent changes), the summary statistics confirm the stylized facts of a mean close to zero with a large standard deviation, and evidence of skewness and excess kurtosis. Evidence from the autocorrelation test for the first 10 lags suggests the existence of mild autocorrelation in all of the rates in first difference but strong autocorrelation in each of the forward premium and the option premium. These results indicate that, while changes in exchange rates, forward rates and option equivalent rates are nearly white noise, the forward premium and the option premium are highly persistent processes (e.g. Backus, Gregory and Telmer, 1993).

Several different unit root tests were conducted to shed light on the integration properties of the time

[14] The summary statistics discussed in this sub-section are not reported to conserve space but they are available from the authors upon request.

series under investigation–our tests include the augmented Dickey-Fuller (ADF) test, the Phillips-Perron test, the KPSS and the nonparametric Breitung test statistic. In keeping with our economic intuition and with the large number of studies of unit root behavior for FX time series, we were in each case unable to reject the unit root null hypothesis for s, f and o, at conventional nominal levels of significance for the level series. On the other hand, differencing the series did appear to induce stationarity in each case. Hence, the unit root tests clearly indicate that spot, forward and option equivalent rates time series are realization from stochastic processes integrated of order one.

4.2 Fama regressions

Our next exercise was to estimate the conventional Fama regression, equation (2) for each currency pair and type of forward rate and option equivalent. This would in principle show the existence of a "forward bias," as recorded in much previous research, and address the question whether there is a similar "option bias" when estimating the Fama regression with the option equivalent.

The results, reported in Table 1, are consistent with the existence of both forward and option bias. Panel A presents the estimation results for the conventional forward contract. We observe that the constant term α is close to zero and often statistically insignificant, whereas β, albeit positive except for the case of the yen, is always estimated to be statistically insignificant. The results are somewhat similar for the option equivalent (Panel B of Table 1). The constant terms are, in most cases, small and insignificantly different from zero. The best estimate of β is for the Swiss franc, where we find positive and significant estimates of β, but the magnitude is close to zero. For all other cases, the slope coefficient β is statistically insignificantly different from zero.

Overall, the results in Table 1 suggest that estimation of the Fama regression using our option equivalent measure rejects unbiasedness and indicates the existence of an option bias puzzle (Panel B) that is consistent with the stylized facts leading to the forward bias puzzle (Panel A).[15]

4.3 Cointegration tests

As discussed in Section 2.1, the Fama regression (2) may not be appropriate for testing the unbiasedness hypothesis because endogeneity issues and an omitted risk premium may render these tests uninformative (e.g. Barnhart, McNown and Wallace, 1999). Regression (2) is essentially a test of UIP under the risk-neutral rational-expectations FX market efficiency hypothesis, which is sufficient but not necessary

[15] Asymptotic standard errors were calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals throughout the paper (Newey and West, 1987).

condition for unbiasedness.

In this section we shift our attention to cointegration analysis, by applying several cointegration tests. We begin with the test proposed by Phillips and Loretan (1991), based on a nonlinear least squares (NLS) estimation procedure which accounts for endogeneity of the regressors. This test is particularly straightforward to implement in that it has a known standard asymptotic distribution, allowing us to test the unbiasedness hypothesis as a test that s_{t+k} and f_t^k cointegrate for the case of forward contracts and that s_{t+k} and o_t^k cointegrate for the case of options contracts–for the case of options, the tests are conducted for all the various definitions of the option equivalent o_t^k given in Section 3.2. The formal test of unbiasedness involves testing cointegration and the hypothesis that $\beta = 1$, where now β is a cointegrating parameter, and then testing the hypothesis that the residuals from the cointegration test are white noise.

Our results, reported in Table 2, show ample support in favor of cointegration between the future spot rate and the current forward rate or the current synthetic forward (option equivalent) rate. Panel A and B display the results for the conventional forward and the option equivalent respectively. The results are very similar. The slope coefficient β (which is now a cointegrating parameter) is generally very close to unity for all cases. The formal test that the cointegrating relationship is of the form $[1, -1]$ is generally not rejected. Also, the statistically significant values of the ADF tests on the residuals from the auxiliary regressions support the hypothesis of cointegration (i.e. reject the null hypothesis of no cointegration).

In addition to the Phillips-Loretan tests, we also carry out a more general nonparametric cointegration test, introduced by Bierens (1997a). This test is appealing in the present context and is deemed superior to standard parametric tests since it has been shown to be capable of detecting cointegration when the data generating process is nonlinear.[16] Although it follows the spirit of reduced rank cointegration tests, the Bierens methodology can consistently estimate the number of cointegrating vectors and also test for parametric restrictions on the cointegrating vectors, on the basis of the ordered solutions of a generalized eigenvalue problem. As for verifying the numbers of cointegrating vectors, Bierens shows how to calculate a λ_{min} test, which is analogous to the Johansen trace test, by testing the null of lower against higher numbers of cointegrating vectors. However, Bierens (1997b) considers the λ_{min} test a tentative outcome and suggests a double check on it by presenting a method for estimating the number of cointegrating vectors ($g_m(r_o)$ test).

[16] Presence of nonlinearity in spot-forward models has been argued by several authors; e.g. see Engel and Hamilton (1990), Clarida, Sarno, Taylor and Valente (2003).

The results from performing the Bierens test are reported in Table 3 (Panels A and B for the conventional forward and the option equivalent respectively). These results, which are again similar between the two different derivatives examined, again provide empirical evidence in favor of cointegration between the spot and the lagged (synthetic) forward rate. Notably, the results from the λ_{min} tests and the $g_m(r_o)$ indicate always cointegration at the 5 percent significance level and suggest the existence of a unique cointegrating vector. Further tests that specify the form of the vector spanning the cointegration space by imposing the restriction of $[1, -1]$ indicated that the null hypothesis of a one-to-one cointegrating relationship could not be rejected at conventional significance levels by the relevant trace test.

Overall, both cointegration techniques employed yield the same outcome, providing ample support in favor of a one-to-one cointegrating relationship between the spot and the lagged (synthetic) forward rate. This is an encouraging result given the difficulties that a large empirical literature finds in detecting an exactly proportional cointegrating relationship between spot and forward rates (e.g. Maynard, 2003). However, this is a necessary condition towards establishing FX unbiasedness, albeit not yet sufficient.

4.4 Residuals tests for FX forward unbiasedness

Following Barnhart, McNown and Wallace (1999), as an additional and more stringent test for unbiasedness, we move on to examine the residual correlation of the errors arising from the cointegrating relationship between the spot and the lagged (synthetic) forward. Given the earlier empirical evidence on the existence of a $[1, -1]$ cointegrating vector in the relationship between the spot and the (synthetic) forward, we construct the deviation from UIP as the difference of the lagged (synthetic) forward rate from the spot rate–i.e. we impose the $[1, -1]$ cointegrating vector–thus generating "restricted" cointegrating residuals. We then employ tests of residual serial correlation, regressing the residuals of each series on their own lagged values (including 4 lags for the 4-lag options and forward series), and a test of cross-correlation, where the residuals are regressed on their own lagged values and on lagged values of the other series (employing again four lags for each different series). We then perform a joint significance coefficient restriction (Wald) test for the null hypothesis of no residual correlation, against the alternative that at least one lag is statistically significant.

The results, presented in Table 4 (Panels A and B), are again very similar between the options and the forward case. Indeed, for all three currencies and for both conventional and synthetic forward rates, the relevant F-test cannot reject the null of no residual autocorrelation and no cross correlation at conventional significance levels (with the only exception of the Swiss franc in one case). Overall, the outcome points towards the validity of the unbiasedness hypothesis.[17]

5 Robustness analysis

In this section we report several robustness checks carried out in order to evaluate the sensitivity of the empirical results reported in the previous section. In particular, we assessed the robustness of our results: (a) to the choice of the number of lags employed in the synthetic forward for the case of 1-month contracts, and (b) to the choice of the maturity of the (synthetic) forward contract, switching to 3-month contracts.

5.1 1-month contracts

Given that our synthetic forward contract contains a mixture of different maturities for the same expiry date, we experimented with taking different numbers of lags, corresponding to different maturities. For that we have selected 8 lags, corresponding to a maturity of 2 months, for which we run the same regressions considered in the core analysis. Our analysis focused on the synthetic forward for each currency, on both the original sample and the void of outliers sample. To conserve space, we report results only for one representative exchange rate, namely dollar-sterling (GBP).

The results confirm the similarities between the conventional and the synthetic forward and show no qualitative difference from the case with 4 lags. Namely, unbiasedness was again rejected on the basis of the Fama regression (Table I in Appendix A, Panels A-B), with the estimates of both the constant α and the slope β being virtually the same as in the core analysis, confirming the existence of an option bias.

On the contrary, ample evidence of a cointegrating relationship between the spot and the synthetic forward was suggested by the Phillips-Loretan and the Bierens tests, which indicate the presence of a $[1, -1]$ cointegrating vector for all cases examined (Table II in Appendix A).

Lastly, we performed an autocorrelation test on the "restricted" residuals, generated with the same method as in the core analysis; however, this time, 8 lags were employed for each currency in the own-

[17] We also performed the same tests on the Phillips-Loretan residuals (not reported but available upon request). These results are qualitatively similar to the results reported for the "restricted" residuals. Furthermore, apart from the above core estimations, in the robustness check we also have employed the Engle-Granger cointegration tests results, since Barnhart, McNown and Wallace (1999) used this test as an element of their empirical work. However, it is well known that the Engle-Granger procedure delivers an estimate of β that, although "superconsistent," can suffer from bias in finite samples, and standard errors do not obey known asymptotic distributions. Therefore, testing restrictions on the cointegrating vector would be futile. It is for this reason that we focus our attention on different tests of cointegration, namely the Philips-Loretan test and the Bierens nonparametric test. Nevertheless, employing the Engle-Granger test on our data confirmed the existence of cointegration detected by the other tests, strengthening our results–details of these test results are not reported to conserve space but they are available from the authors upon request.

and cross-correlation tests. The majority of outcomes could not reject the null of no autocorrelation at conventional significance levels (Table III in Appendix A). Nevertheless, there are minor exceptions where serial or cross correlation is detected, but they did not entice any specific pattern. Thus, this evidence notwithstanding, we conclude that our core results are robust to changes in the lags employed and offer support to the unbiasedness hypothesis on the basis of cointegration and residuals tests.

5.2 3-month contracts

We then re-estimated the core regressions for each exchange rate examined using a 3-month forward contract and a 3-month synthetic forward contract, at the weekly frequency, to assess the robustness to the choice of the contract maturity. In order to construct our synthetic contract we have chosen the expiry dates of the fixed quarterly cycles (March, June, September and December), and gathered all relevant contracts. Again, we have 3-month synthetic forwards with a range of maturities from 4 to 52 weeks. For reasons of consistency to the forward case, we choose to work with 12 lags.

The results are, again, very similar between the forward and the synthetic forward case and between the 1-month and the 3-month contracts. The Fama regression for the conventional forward contract presents negative and insignificant coefficients for the β slope parameter, whereas the constant term α is estimated to be close to zero (albeit significant for the yen and the Swiss franc)–Panel A of Table I in Appendix B. These results are comparable with the ones obtained from estimating the Fama regression with the option premium (equation (7)), for which we report the results only for GBP, as a representative rate–Panel B of Table I in Appendix B.

Shifting our attention to the cointegration tests, again both the Phillips-Loretan test and the Bierens test detect the existence of a $[1, -1]$ cointegrating relationship in the case of the spot-forward as well as spot-option cases (Table II in Appendix B). Finally, the autocorrelation and cross-correlation tests on the residuals (this time performed with 12 lags) lead to similar conclusions as in the core analysis–i.e. the null of no autocorrelation cannot be rejected (Table III in Appendix B).[18]

6 Conclusions

Armed with several tests proposed by the literature testing forward rate unbiasedness in the FX market, this paper provides a simple, yet intuitive bridge to a different derivatives market, the currency options market, as a vehicle of forming expectations about future spot exchange rates. Our main focus is on

[18] Again, we report results only for GBP as a representative exchange rate to conserve space.

performing tests of the unbiasedness hypothesis. To that end, we used the conventional forward rate and also introduced an option equivalent (synthetic forward) contract. We then apply some prominent tests of unbiasedness, as suggested by the relevant literature. We adopt tests based on the standard UIP condition in a stationary setting as well as cointegration tests for unbiasedness of the (synthetic) forward rate, the latter combined with residual autocorrelation tests.

Our research provides encouraging results. We manage to bridge the distance between the forward (OTC) market and the options (exchange traded) market, by directly comparing the test results obtained for the two markets. Viewed from a different angle, our research offers a novel robustness check to the tenacity of the well-documented forward bias anomaly that characterizes the relevant literature.

We record no qualitative difference between the two types of derivatives products in our results. Specifically, our results suggest the existence of an "options bias," similar to the forward bias, frequently recorded in the relevant literature estimating stationary regressions of the exchange rate change on the lagged forward premium. This finding indicates, in turn, violation of market efficiency in its risk neutral formulation as implied by UIP, possibly as a consequence of the existence of a risk premium. We therefore shift our attention to a cointegration and a residual correlation analysis that allows for the endogeneity problems caused by a potential unobserved risk premium term. This analysis attests that indeed the (synthetic) forward is an unbiased predictor of the future spot rate.

Overall we find that, if one is willing to entertain the possibility that there is a non-zero foreign exchange risk premium, then forward and options provide optimal predictions consistent with the notion of unbiasedness.

Table 1. Fama regressions: 1-month contracts (k=4)

Panel A) Forward premium Fama regressions

	α	SE(α)	β	SE(β)
GBP	0.002*	(0.001)	0.371	(0.292)
JPY	0.003**	(0.001)	-0.260	(0.248)
CHF	0.002**	(0.001)	0.148	(0.274)

Panel B) Option premium Fama regressions

		original sample				outliers removed			
		α	SE(α)	β	SE(β)	α	SE(α)	β	SE(β)
GBP	ATM	0.001	(0.001)	0.060	(0.065)	0.001	(0.001)	0.088	(0.070)
	Average	0.001	(0.001)	0.021	(0.065)	0.001	(0.001)	0.027	(0.078)
	Last Trade	0.001	(0.001)	-0.039	(0.050)	0.001	(0.001)	-0.040	(0.060)
	Median	0.001	(0.001)	0.001	(0.057)	0.001	(0.001)	0.024	(0.067)
	W. average	0.001	(0.000)	-0.011	(0.056)	0.001	(0.001)	-0.033	(0.069)
JPY	ATM	0.003**	(0.001)	-0.009	(0.047)	0.002**	(0.001)	-0.126	(0.110)
	Average	0.003**	(0.001)	-0.031	(0.047)	0.002**	(0.001)	-0.197*	(0.080)
	Last Trade	0.003**	(0.001)	-0.023	(0.034)	0.003**	(0.001)	-0.082	(0.052)
	Median	0.003**	(0.001)	-0.016	(0.040)	0.003**	(0.001)	-0.094	(0.064)
	W. average	0.003**	(0.001)	-0.028	(0.042)	0.002**	(0.001)	-0.107***	(0.064)
CHF	ATM	0.002	(0.001)	0.135*	(0.065)	0.002	(0.001)	0.165**	(0.075)
	Average	0.002	(0.001)	0.117	(0.064)	0.002	(0.001)	0.262*	(0.080)
	Last Trade	0.002	(0.001)	0.060	(0.049)	0.002	(0.001)	0.132*	(0.052)
	Median	0.002	(0.001)	0.089	(0.050)	0.002	(0.001)	0.063*	(3.420)
	W. average	0.002*	(0.001)	0.044	(0.052)	0.002	(0.001)	0.068**	(2.045)

Notes. *Panel A)* The table shows the results from estimating, by ordinary least squares, the conventional forward premium (Fama) regression in equation (2). *Panel B)* The table shows the results from estimating, by ordinary least squares, the option premium (Fama) regression in equation (7). For both panels, figures in parentheses (SE(α) and SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals up to the third decimal point (Newey and West, 1987). One and two asterisks denote statistical significance at the 5 and 1 percent level respectively.

Table 2. Phillips-Loretan cointegration tests: 1-month contracts (k=4)

Panel A) Spot-forward relationship

	β	SE(β)	ADF	F
GBP	1.003**	(0.003)	-15.268**	[0.236]
JPY	0.993**	(0.007)	-15.908**	[0.318]
CHF	1.011**	(0.040)	-16.031**	[0.792]

Panel B) Spot-option relationship

		original sample				outliers removed			
		β	SE(β)	ADF	F	β	SE(β)	ADF	F
GBP	ATM	1.006**	(0.012)	-15.562**	[0.563]	1.007**	(0.015)	-15.585**	0.222
	Average	1.025**	(0.093)	-15.739**	[0.531]	1.021**	(0.073)	-15.702**	0.080
	Last Trade	0.978**	(0.101)	-15.759**	[0.720]	1.123	(2.555)	-15.593**	0.002
	Median	1.125	(2.206)	-15.773**	[0.793]	1.025**	(0.118)	-15.853**	0.046
	W. average	0.972**	(0.124)	-15.979**	[0.782]	0.988**	(0.040)	-15.890**	0.097
JPY	ATM	0.988**	(0.017)	-15.200**	[0.660]	0.997**	(0.005)	-15.713**	0.391
	Average	0.991**	(0.012)	-15.117**	[0.650]	0.998**	(0.004)	-15.584**	0.574
	Last Trade	0.990**	(0.013)	-15.064**	[0.653]	0.996**	(0.006)	-15.409**	0.445
	Median	0.987**	(0.016)	-14.857**	[0.745]	0.995**	(0.008)	-15.434**	0.805
	W. average	0.989**	(0.013)	-14.890**	[0.691]	0.996**	(0.006)	-15.278**	0.380
CHF	ATM	1.008**	(0.011)	-15.466**	[0.650]	1.008**	(0.011)	-15.608**	0.533
	Average	1.003**	(0.006)	-15.491**	[0.908]	1.003**	(0.006)	-15.650**	0.369
	Last Trade	1.010**	(0.016)	-15.516**	[0.829]	1.010**	(0.016)	-15.445**	0.421
	Median	1.003**	(0.007)	-15.489**	[0.919]	1.004**	(0.007)	-15.874**	0.270
	W. Average	1.012**	(0.018)	-15.585**	[0.723]	1.012**	(0.018)	-15.666**	0.430

Notes. *Panel A)* The table presents the results from testing for cointegration between the spot rate, s_{t+4} and the forward rate, f_t^4 using the Phillips-Loretan (1991) test. *Panel B)* The table presents the results from testing for cointegration between the spot rate, s_{t+4} and the synthetic forward, o_{t,j_h}^4 rate using the Phillips-Loretan (1991) test, where j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. *For both panels*, β denotes the cointegrating parameter and figures in parentheses (SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals (Newey and West, 1987). ADF is the Augmented Dickey Fuller test statistic for a unit root in the residuals (i.e. for no cointegration). The column F gives the p-value from the relevant F-statistic for the null hypothesis that the cointegrating vector is $[1, -1]$. One and two asterisks denote statistical significance at the 5 and 1 percent level respectively.

Table 3. Bierens nonparametric cointegration tests: 1-month contracts (k=4)

Panel A) Spot-forward relationship

	$\lambda_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
GBP	[0.000]*	7.30×10^7	1.020
	[0.078]	$\mathbf{1.96 \times 10^0}$	
		1.02×10^4	
JPY	[0.000]*	3.68×10^5	1.090
	[0.423]	$\mathbf{1.31 \times 10^1}$	
		2.02×10^6	
CHF	[0.000]*	5.61×10^4	1.370
	[0.137]	$\mathbf{8.18 \times 10^2}$	
		1.33×10^7	

(continued ...)

(... Table 3 continued)

Panel B) Spot-option relationship

GBP)

	original sample			outliers removed		
	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0,1,2$	Trace test	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0,1,2$	Trace test
ATM	[0.000]*	3.21×10^{1}	1.010	[0.000]*	3.89×10^{0}	1.010
	[0.062]	**7.20×10^{-3}**		[0.062]	**5.88×10^{-3}**	
		2.39×10^{1}			1.97×10^{1}	
Average	[0.000]*	7.17×10^{9}	1.140	[0.000]*	8.74×10^{13}	1.050
	[0.058]	**3.62×10^{-2}**		[0.061]	**2.72×10^{-6}**	
		1.09×10^{2}			8.79×10^{-3}	
Last Trade	[0.000]*	4.38×10^{1}	1.010	[0.000]*	8.36×10^{11}	1.010
	[0.064]	**4.90×10^{-3}**		[0.063]	**2.68×10^{-4}**	
		1.75×10^{1}			9.18×10^{-1}	
Median	[0.000]*	2.31×10^{12}	1.080	[0.000]*	1.37×10^{12}	1.050
	[0.059]	**1.08×10^{-4}**		0.059	**1.85×10^{-4}**	
		3.33×10^{-1}			5.59×10^{-1}	
W. Average	[0.000]*	1.34×10^{12}	1.040	[0.000]*	5.43×10^{11}	1.020
	[0.061]	**1.77×10^{-4}**		[0.061]	**4.36×10^{-4}**	
		5.73×10^{-1}			1.41×10^{0}	

JPY)

	original sample			outliers removed		
	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0,1,2$	Trace test	λ^a_{min} T1, T2	$g_m(r_o)$ $r_o = 0,1,2$	Trace test
ATM	[0.000]*	1.07×10^{11}	1.040	[0.000]*	2.08×10^{7}	1.100
	[0.370]	**5.94×10^{-5}**	1.040	[0.377]	**2.93×10^{-1}**	
		7.05×10^{0}			3.62×10^{4}	
Average	[0.000]*	3.28×10^{10}	1.180	[0.000]*	6.47×10^{7}	1.190
	[0.368]	**1.96×10^{-4}**	1.180	[0.380]	**9.50×10^{-2}**	
		2.30×10^{1}			1.17×10^{4}	
Last Trade	[0.000]*	3.77×10^{9}	1.020	[0.000]*	1.88×10^{7}	1.180
	[0.378]	**1.61×10^{-3}**	1.020	[0.379]	**3.21×10^{-1}**	
		2.00×10^{2}			4.02×10^{4}	
Median	[0.000]*	6.11×10^{8}	1.550	[0.000]*	1.99×10^{8}	1.270
	0.362	**1.08×10^{-02}**	1.550	0.370	**3.19×10^{-2}**	
		1.23×10^{3}			3.80×10^{3}	
W. Average	[0.000]*	1.96×10^{10}	1.190	[0.000]*	7.12×10^{7}	1.290
	[0.364]	**3.33×10^{-04}**	1.190	[0.371]	**8.88×10^{-2}**	
		3.84E $\times 10^{1}$			1.06×10^{4}	

(continued ...)

(... Table 3 continued)

CHF)

	original sample			outliers removed		
	$\lambda^a_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test	$\lambda^a_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
ATM	[0.000]*	2.55×10^7	1.180	[0.000]*	9.92×10^6	1.210
	[0.120]	$\mathbf{2.39 \times 10^0}$		[0.121]	$\mathbf{6.03 \times 10^0}$	
		3.01×10^4			7.73×10^4	
Average	[0.000]*	5.49×10^7	1.150	[0.000]*	1.29×10^7	1.170
	[0.120]	$\mathbf{1.12 \times 10^0}$		[0.118]	$\mathbf{4.85 \times 10^0}$	
		1.40×10^4			5.97×10^4	
Last Trade	[0.000]*	5.38×10^7	1.160	[0.000]*	1.92×10^7	1.170
	[0.123]	$\mathbf{1.07 \times 10^0}$		[0.121]	$\mathbf{3.09 \times 10^0}$	
		1.43×10^4			3.99×10^4	
Median	[0.000]*	3.05×10^7	1.170	[0.000]*	8.68×10^6	1.170
	[0.120]	$\mathbf{1.99 \times 10^0}$		[0.118]	$\mathbf{7.28 \times 10^0}$	
		2.51×10^4			8.84×10^4	
W. Average	[0.000]*	1.06×10^8	1.130	[0.000]*	2.08×10^7	1.150
	[0.119]	$\mathbf{5.79 \times 10^1}$		[0.118]	$\mathbf{3.04 \times 10^0}$	
		7.22×10^3			3.69×10^4	

Notes. The tables present the results from the nonparametric cointegration tests of Bierens (1997a) applied to the spot-forward relationship (s_{t+4} and f^4_t) and the spot-option (s_{t+4} and o^4_{t,j_h}) relationship; j = *ATM, Average, Last Trade, Median and W. Average*. The first column of results ($\lambda_{\min}$) shows the p-values of the $\lambda_{\min}$ test statistic for T1 (which is H_o: *r=0 vs.* H_1 *: r=1*) and for T2 (which is H_o: *r=1 vs.* H_1 *r=2*) respectively. The second column calculates $r_m = \arg\min_{r_o \leq 2}\{g_m(r_o)\}$ for $m = 2$, where r_o is the number of cointegrating vectors; the table presents the r_m values for $r_o = 0, 1, 2$. The number in bold emphasizes the minimum r_m value, which indicates the number of cointegrating relationships identified by the Bierens test. The final column presents the results from the trace test for the null hypothesis that the cointegrating vector is $[1, -1]$, i.e. $H_o : \beta' = (1, -1)$; the appropriate 5-percent critical value is 4.70. The asterisk denotes statistical significance at the 5 percent level.

Table 4. Residual tests: 1-month contracts (k=4)

Panel A) Residual correlation tests for spot-forward

	AC	CC
GBP	[0.399]	[0.862]
JPY	[0.327]	[0.322]
CHF	[0.178]	[0.543]

Panel B) Residual correlation tests for spot-options

		original sample		outliers removed	
		AC	CC	AC	CC
GBP	ATM	[0.540]	[0.836]	[0.610]	[0.957]
	Average	[0.602]	[0.790]	[0.764]	[0.831]
	Last Trade	[0.755]	[0.827]	[0.445]	[0.852]
	Median	[0.534]	[0.793]	[0.663]	[0.903]
	W. average	[0.530]	[0.769]	[0.787]	[0.855]
JPY	ATM	[0.087]	[0.369]	[0.063]	[0.128]
	Average	[0.119]	[0.165]	[0.079]	[0.279]
	Last Trade	[0.179]	[0.073]	[0.318]	[0.187]
	Median	[0.143]	[0.082]	[0.271]	[0.151]
	W. average	[0.148]	[0.126]	[0.055]	[0.166]
CHF	ATM	[0.194]	[0.265]	[0.436]	[0.688]
	Average	[0.066]	[0.132]	[0.352]	[0.570]
	Last Trade	[0.167]	[0.285]	[0.447]	[0.485]
	Median	[0.062]	[0.072]	[0.148]	[0.177]
	W. Average	[0.019]*	[0.026]*	[0.247]	[0.179]

Notes. The tables present the p-values for the relevant F-statistics for joint coefficient restriction on $s_{t+k} - f_t$ for the forward and $s_{t+k} - o_{t,j}$ for the synthetic forward, where j = *ATM, Average, Last Trade, Median and W. Average*. We perform tests for autocorrelation (AC) and cross correlation (CC) for the spot-forward and the spot-option relationships. For the case of the forward, the F-statistic is distributed with (4, 925) degrees of freedom for the AC test and (12, 917) degrees of freedom for the CC test respectively. For the case of options the F-statistic is distributed with (4, 922) degrees of freedom for the AC test and (12, 914) degrees of freedom for the CC test respectively.

A Appendix: Robustness results: 1-month contracts

Table I. Fama regression on the spot-option relationship: GBP (1-month contracts, k=8)

	Original sample				Outliers removed			
	α	SE(α)	β	SE(β)	α	SE(α)	β	SE(β)
ATM	0.002	(0.002)	0.012	(0.091)	0.002	(0.001)	0.038	(0.101)
Average	0.002	(0.001)	-0.034	(0.091)	0.002	(0.001)	-0.028	(0.110)
Last Trade	0.002	(0.001)	-0.069	(0.069)	0.002	(0.001)	-0.022	(0.084)
Median	0.002	(0.001)	-0.065	(0.079)	0.002	(0.001)	-0.055	(0.093)
W. average	0.002	(0.001)	-0.101	(0.078)	0.002	(0.001)	-0.129	(0.097)

Notes. The table shows the results from estimating, by ordinary least squares, the option premium (Fama) regression in equation (7). Figures in parentheses (SE(α) and SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals up to the third decimal point (Newey and West, 1987).

Table II. Cointegration tests: GBP (1-mont contracts, k=8)

Panel A) Phillips-Loretan tests on the spot-option relationship

	original sample				outliers removed			
	β	SE(β)	ADF	F	β	SE(β)	ADF	F
ATM	1.008**	(0.019)	-9.163**	[0.638]	1.012**	(0.033)	-8.766**	[0.721]
Average	1.089	(0.968)	-9.193**	[0.926]	1.058**	(0.093)	-9.184**	[0.907]
Last Trade	1.048**	(0.304)	-9.061**	[0.877]	1.012**	(0.024)	-9.189**	[0.628]
Median	1.002**	(0.019)	-9.063**	[0.940]	1.003**	(0.016)	-9.034**	[0.939]
W. average	0.990**	(0.025)	-9.415**	[0.698]	0.993**	(0.018)	-9.380**	[0.711]

(continued ...)

(... Table II continued)

Panel B) Bierens nonparametric tests on the GBP spot-option relationship

	original sample			outliers removed		
	$\lambda^a_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test	$\lambda^a_{\min}$ T1, T2	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace test
ATM	[0.000]*	7.25×10^8	1.000	[0.000]*	7.74×10^8	1.010
	[0.066]	$\mathbf{2.75 \times 10^{-1}}$		[0.067]	$\mathbf{2.51 \times 10^{-1}}$	
		1.04×10^3			9.75×10^2	
Average	[0.000]*	2.45×10^9	1.120	[0.000]*	3.30×10^9	1.010
	[0.064]	$\mathbf{8.60 \times 10^{-2}}$		[0.064]	$\mathbf{6.51 \times 10^{-1}}$	
		3.07×10^2			2.29×10^1	
Last Trade	[0.000]*	7.90×10^8	1.000	[0.000]*	1.93×10^9	1.000
	[0.066]	$\mathbf{2.53 \times 10^{-1}}$		[0.065]	$\mathbf{1.08 \times 10^{-1}}$	
		9.55×10^2			3.91×10^1	
Median	[0.000]*	1.92×10^9	1.020	[0.000]*	1.74×10^9	1.010
	[0.062]	$\mathbf{1.12 \times 10^{-1}}$		[0.061]	$\mathbf{1.34 \times 10^{-1}}$	
		3.92×10^2			4.34×10^2	
W. Average	[0.000]*	1.91×10^9	1.010	[0.000]*	1.53×10^9	1.010
	[0.063]	$\mathbf{1.15 \times 10^{-1}}$		[0.063]	$\mathbf{1.43 \times 10^{-1}}$	
		3.95×10^1			4.93×10^2	

Notes. *Panel A)* The table presents the results from testing for cointegration between the spot rate, s_{t+8} and the synthetic forward, o^8_{t,j_h} rate using the Phillips-Loretan (1991) test, where $j = ATM, Average,$ *Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. β denotes the cointegrating parameter and figures in parentheses (SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals (Newey and West, 1987). ADF is the Augmented Dickey Fuller test statistic for a unit root in the residuals (i.e. for no cointegration). The column F gives the p-value from the relevant F-statistic for the null hypothesis that the cointegrating vector is $[1, -1]$. One and two asterisks denote statistical significance at the 5 and 1 percent level respectively.

Panel B) The tables present the results from the nonparametric cointegration tests of Bierens (1997a) applied to the spot-option (s_{t+8} and o^8_{t,j_h}) relationship. The first column of results ($\lambda_{\min}$) shows the p-values of the $\lambda_{\min}$ test statistic for T1 (which is H_o: *r=0 vs.* H_1 *: r=1*) and for T2 (which is H_o: *r=1 vs.* H_1 *r=2*) respectively. The second column calculates $r_m = \arg\min_{r_o \leq 2}\{g_m(r_o)\}$ for $m = 2$, where r_o is the number of cointegrating vectors; the table presents the r_m values for $r_o = 0, 1, 2$. The number in bold emphasizes the minimum r_m value, which indicates the number of cointegrating relationships identified by the Bierens test. The final column presents the results from the trace test for the null hypothesis that the cointegrating vector is $[1, -1]$, i.e. $H_o : \beta' = (1, -1)$; the appropriate 5-percent critical value is 4.70. The asterisk denotes statistical significance at the 5 percent level.

Table III. Residual correlations test: GBP (1-month contracts, k=8)

	original sample		outliers removed	
	AC	CC	AC	CC
ATM	[0.998]	[0.966]	[0.636]	[0.392]
Average	[0.978]	[0.891]	[0.989]	[0.839]
Last Trade	[0.971]	[0.995]	[0.791]	[0.953]
Median	[0.935]	[0.415]	[0.935]	[0.538]
W. average	[0.808]	[0.953]	[0.842]	[0.944]

Notes. The table presents the p-values for the relevant F-statistics for joint coefficient restriction on $s_{t+k} - f_t$ for the forward and $s_{t+k} - o_{t,j}$ for the synthetic forward; j = *ATM, Average, Last Trade, Median and W. Average.* We perform tests for autocorrelation (AC) and cross correlation (CC) for the spot-option relationships. The F-statistic is distributed with (8, 910) degrees of freedom for the AC test and (24, 897) degrees of freedom for the CC test respectively.

B Appendix: Robustness results: 3-month contracts

Table I. Fama regressions: 3-month contracts (k=12)

Panel A) Forward premium Fama regressions

	α	SE(α)	β	SE(β)
GBP	0.000	(0.002)	-0.309	(0.081)
JPY	0.025*	(0.003)	-2.359*	(0.329)
CHF	0.010*	(0.002)	-1.056*	(0.300)

Panel B) Option premium Fama regressions: GBP

	Original sample				Outliers removed			
	α	SE(α)	β	SE(β)	α	SE(α)	β	SE(β)
ATM	0.003	(0.002)	0.131*	(0.058)	0.002	(0.002)	0.110	(0.075)
Average	0.002	(0.002)	0.046	(0.061)	0.003	(0.002)	0.141	(0.086)
Last Trade	0.002	(0.002)	0.001	(0.048)	0.003	(0.002)	0.093	(0.068)
Mean	0.002	(0.002)	0.022	(0.051)	0.002	(0.002)	0.102	(0.071)
W. average	0.002	(0.002)	0.049	(0.054)	0.003	(0.002)	0.115	(0.076)

Notes. *Panel A)* The table shows the results from estimating, by ordinary least squares, the conventional forward premium (Fama) regression in equation (2). *Panel B)* The table shows the results from estimating, by ordinary least squares, the option premium (Fama) regression in equation (7) applied to the data for dollar-sterling. For both panels, figures in parentheses (SE(α) and SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals up to the third decimal point (Newey and West, 1987). One and two asterisks denote statistical significance at the 5 and 1 percent level respectively.

Table II: Cointegration tests: 3-month contracts (k=12)

Panel A1) Phillips-Loretan cointegration tests for the spot-forward relationship

	β	SE(β)	ADF	$\mathcal{F}$
GBP	0.984*	0.251	-6.668*	0.004
JPY	0.991*	0.001	-6.596*	110.589*
CHF	0.993*	0.0017	-6.601*	15.087*

Panel A2) Phillips-Loretan cointegration tests for the spot-option relationship: GBP

	original sample				outliers removed			
	β	SE(β)	ADF	$\mathcal{F}$	β	SE(β)	ADF	$\mathcal{F}$
ATM	1.006**	(0.010)	-7.217**	0.444	1.005**	(0.011)	-7.126**	0.204
Average	1.010	(0.203)	-7.251**	0.256	1.006**	(0.009)	-7.172**	0.442
Last Trade	1.013**	(0.034)	-7.536**	0.145	1.006**	(0.012)	-7.381**	0.282
Mean	1.008**	(0.022)	-7.395**	0.129	1.004**	(0.010)	-7.312**	0.159
W. average	1.010**	(0.020)	-7.482**	0.255	1.005**	(0.010)	-7.414**	0.256

Panel B1) Bierens nonparametric cointegration tests for the spot-forward relationship

	$\lambda^a_{\min}$ *T1, T2*	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace Test
GBP	[0.000]*	7.30×10^7	1.020
	[0.078]	$\mathbf{1.96 \times 10^0}$	
		1.02×10^4	
JPY	[0.000]*	3.68×10^5	1.090
	[0.423]	$\mathbf{1.31 \times 10^1}$	
		2.02×10^6	
CHF	[0.000]*	5.61×10^4	1.370
	[0.137]	$\mathbf{8.18 \times 10^2}$	
		1.33×10^7	

(continued ...)

(...Table II continued)

Panel B2) Bierens nonparametric cointegration tests for the spot-option relationship: GBP

	original sample			outliers removed		
	$\lambda^a_{\min}$ *T1, T2*	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace Test	$\lambda^a_{\min}$ *T1, T2*	$g_m(r_o)$ $r_o = 0, 1, 2$	Trace Test
ATM	[0.000]*	1.33×10^{9}	1.000	[0.000]*	1.40×10^{8}	1.000
	[0.095]	$\mathbf{7.14 \times 10^{-2}}$		[0.093]	$\mathbf{7.11 \times 10^{-1}}$	
		5.49×10^{2}			5.20×10^{3}	
Average	[0.000]*	5.78×10^{0}	1.010	[0.000]*	1.41×10^{9}	1.010
	[0.089]	$\mathbf{1.86 \times 10^{-3}}$		[0.089]	$\mathbf{7.67 \times 10^{-2}}$	
		1.26×10^{1}			5.15×10^{2}	
Last Trade	[0.000]*	3.27×10^{0}	1.010	[0.000]*	2.75×10^{10}	1.010
	[0.092]	$\mathbf{3.08 \times 10^{-3}}$		[0.077]	$\mathbf{5.34 \times 10^{-3}}$	
		2.23×10^{1}			2.79×10^{1}	
Mean	[0.000]*	3.87×10^{13}	1.010	[0.000]*	6.11×10^{8}	1.010
	[0.089]	$\mathbf{2.76 \times 10^{-6}}$		[0.090]	$\mathbf{1.08 \times 10^{2}}$	
		1.88×10^{-2}			1.23×10^{3}	
W. Average	[0.000]*	8.55×10^{10}	1.010	[0.000]*	8.55×10^{10}	1.010
	[0.086]	$\mathbf{1.37 \times 10^{-3}}$		[0.086]	$\mathbf{1.37 \times 10^{-3}}$	
		8.53×10^{0}			8.53×10^{0}	

Notes. *Panels A1) and A2)* Panel A1 presents the results from testing for cointegration between the spot rate, s_{t+12} and the forward rate, f_t^{12} using the Phillips-Loretan (1991) test. Panel A2 presents the results from testing for cointegration between the spot rate, s_{t+12} and the synthetic forward, o^{12}_{t,j_h} rate using the Phillips-Loretan (1991) test, where j = *ATM, Average, Last Trade, Median and W. Average* corresponds to the five different methods of data construction described in Section 3.2; and the subscript $h = a, or$ stands for the analysis of the original series (a) and the series after the removal of the outliers (or) respectively. Results are for the dollar-sterling exchange rate. For both panels, β denotes the cointegrating parameter and figures in parentheses (SE(β)) are asymptotic standard errors calculated using an autocorrelation and heteroskedasticity consistent matrix of residuals (Newey and West, 1987). ADF is the Augmented Dickey Fuller test statistic for a unit root in the residuals (i.e. for no cointegration). The column F gives the p-value from the relevant F-statistic for the null hypothesis that the cointegrating vector is $[1, -1]$. One and two asterisks denote statistical significance at the 5 and 1 percent level respectively.

Panels B1) and B2) The tables present the results from the nonparametric cointegration tests of Bierens (1997a) applied to the spot-forward relationship (s_{t+12} and f_t^{12}), reported in Panel B1, and the spot-option relationship (s_{t+12} and o^{12}_{t,j_h}), reported in Panel B2. The first column of results ($\lambda_{\min}$) shows the p-values of the $\lambda_{\min}$ test statistic for T1 (which is H_o: *r=0 vs. H_1: r=1*) and for T2 (which is H_o: *r=1 vs. H_1 r=2*) respectively. The second column calculates $r_m = \arg\min_{r_o \leq 2}\{g_m(r_o)\}$ for $m = 2$, where r_o is the number of cointegrating vectors; the table presents the r_m values for $r_o = 0, 1, 2$. The number in bold emphasizes the minimum r_m value, which indicates the number of cointegrating relationships identified by the Bierens test. The final column presents the results from the trace test for the null hypothesis that the cointegrating vector is $[1, -1]$, i.e. $H_o : \beta' = (1, -1)$; the appropriate 5-percent critical value is 4.70. The asterisk denotes statistical significance at the 5 percent level.

Table III. Residual tests: 3 month contracts (k=12)

Panel A) Residual correlation tests for spot-forward

		original sample	
		AC	CC
GBP	ATM	[0.866]	[0.340]
JPY	Average	[0.218]	[0.391]
CHF	Last Trade	[0.709]	[0.567]

Panel B) Residual correlation tests for spot-options

	original sample		outliers removed	
	AC	CC	AC	CC
ATM	[0.866]	[0.340]	[0.101]	[0.101]
Average	[0.218]	[0.391]	[0.187]	[0.448]
Last Trade	[0.709]	[0.567]	[0.496]	[0.488]
Mean	[0.556]	[0.858]	[0.567]	[0.819]
W. average	[0.207]	[0.509]	[0.207]	[0.638]

Notes. The tables present the p-values for the relevant F-statistics for joint coefficient restriction on $s_{t+k} - f_t$ for the forward and $s_{t+k} - o_{t,j}$ for the synthetic forward; j = *ATM, Average, Last Trade, Median and W. Average*. We perform tests for autocorrelation (AC) and cross correlation (CC) for the spot-forward and the spot-option relationships. For the case of the forward, the F-statistic is distributed with (12, 897) degrees of freedom for the AC test and (36, 873) degrees of freedom for the CC test respectively. For the case of options the F-statistic is distributed with (12, 892) degrees of freedom for the AC test and (36, 867) degrees of freedom for the CC test respectively.

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