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## Interest parity conditions during the classical gold standard (1880 -1914) - Evidence from the investment demand for bills of exchange in Europe

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# Interest parity conditions during the classical gold standard (1880 -1914) - Evidence from the investment demand for bills of exchange in Europe

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**Abstract**

This paper examines several versions of the (covered and uncovered) interest parity condition that arguably held as regards the investment demand for bills of exchange during the classical gold standard (1880 - 1914). Contemporaneous guide books about the foreign exchanges report that close connections between the exchange and interest (or discount) rates arose mainly between London and the major financial centres on the European continent. As implied by the interest parity condition, and in particular when future exchange rate movements were covered via a suitable long-bill transaction, weekly data suggest indeed that between Paris, Amsterdam, Berlin and London, the return from discounting bills of exchange in the local money market was roughly equivalent to the (exchange rate adjusted) return from investing in foreign bills.

*JEL classification: F31, N13, N23*

*Keywords: Covered interest parity condition; Exchange rates; Gold standard; Uncovered interest parity condition*

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# 1 Introduction

> "The Foreign Exchanges [...] are the barometer of the Money Market. Between the price of London bills, as expressed in the current rate of discount, and the price of foreign bills, as expressed in the current exchange rate, there exists at times a close sympathy [...]." Clare (1902, p.87)

Thanks to the (covered and uncovered) interest parity condition, the connections between the foreign exchange and money markets feature nowadays prominently in international finance. It is, however, remarkable that the introductory quote has been taken from a practical guide book first published back in 1891 to provide "A Money-Market Primer and Key to the Exchanges" bearing in mind that, at the time, research on the foreign exchanges was still in its infancy. Arguably, the first hints at the interest parity condition appeared in Thornton (1802), whilst Goschen (1861) and Lotz (1889) published more extensive, but still rather informal, discussions. It was only during the first part of the $20^{th}$ century when the theory that exchange rate movements offset, on average, the differences in interest rates between currencies had definitely established itself in the academic literature (Einzig, 1962, ch.17 and 21). Path-breaking contributions to the uncovered and covered interest parity condition are usually ascribed to, respectively, Fisher (1896, ch.9) and Keynes (1923, ch.3.4). Conversely, as reflected by the introductory quote, among bankers, an informal understanding of the connections between the exchange and interest rate—or at the time the discount rate—must already have existed during the second half of the $19^{th}$ century. Einzig (1962, pp.184-185) reports indeed that at the time, "the establishment of closer relations between banks [...] led to a considerable increase in the volume of exchange arbitrage [...] and also the volume of interest arbitrage" and that the corresponding techniques were "described in detail in innumerable practical books and articles". In particular, it was arguably mainly in the wealthier parts of Europe where banks had the option of reinvesting idle deposits in the local money market, or earning short-term interest by purchasing bills of exchange issued on foreign financial centres. Reflecting the importance of this so-called "investment demand for bills of exchange", it seems to have been well-known that e.g. "bankers of the Continent begin buying up London paper as soon as the interest obtainable from it exceeds the rate to be earned on native acceptances" (Clare, 1895, p.93). Similar guide book discussions about interest parity transactions with foreign bills of exchange can be found in Clare (1902, ch.9), Spalding (1915, ch.7), and Thomas (1921, ch.8). In his famous portrait of the London money market in Lombard Street, even Bagehot (1873, pp.45-46) touches briefly on some aspects of this issue.

It is perhaps surprising that the large literature on the empirical aspects of the interest parity condition has by and large ignored the possible connections between interest and exchange rates at the end of the $19^{th}$ and the beginning of the $20^{th}$ century. An exception is Lothian and Wu (2011), who have looked at time series covering the past 200 years. However, due to the annual frequency of their data, they only check whether the uncovered interest parity (UIP) condition holds over the very long haul. Furthermore, Goodhart (1969) and more recently Coleman (2012) have estimated some UIP-regressions for the decades around 1900. Yet, their focus lay on the New York money market and, more specifically, its characteristic seasonal exchange and interest rate fluctuations reflecting annually recurring cycles in agricultural activity. Though international capital flows could have exploited these seasonal patters, Coleman (2012) suggests that interest parity transactions involving the dollar-sterling exchange rate did frequently not occur (though when they occurred, the UIP-condition held relatively well). Among other things, this result is attributed to the substantial transaction costs of taking speculative positions in the New York sterling-bill market. Conversely, during the classical gold standard, the London market for bills of exchange was larger, more liquid, and seems to have been much closer integrated with the financial centres across the English Channel. Concurring with this, the contemporaneous guide books emphasise that, rather than with New York, a nexus existed between the discount rate in the open money markets of Paris, Berlin, or Amsterdam, and the demand for London bills of exchange (Clare, 1902, pp.129ff.). Finally, turning to the covered interest parity (CIP) condition, Flandreau and Komlos provide an empirical study on early arbitrage transactions between Berlin and Vienna, where a forward foreign exchange market had already developed during the second half of the $19^{th}$ century. However, this seems to have been a fairly isolated event in the sense that for the major currencies around the world, regular forward exchange transactions did no occur until after the First World War (Einzig, 1962, pp.182-183; Einzig, 1967, pp.7-8).

Against this background, this paper endeavours to contribute to the literature by testing, via modern econometric techniques, how closely the interest (or discount) and exchange rates were connected between the leading European financial centres (London, Paris, Berlin, Amsterdam) during the classical gold standard (1880 - 1914). Thereby, the fact that bills of exchange were the dominant security for short-term investments as well as the role of London as financial hub will be taken into account. The results suggest that the implications of the interest parity condition are often supported by the (weekly) data covering the major European currencies backed by gold. Specifically, when comparing the discount rate in the local money market with the exchange rate adjusted return in a foreign financial centre, the postulated proportional relationship tends to arise when future exchange rate movements are covered by means of a suitable long-bill investment. Furthermore, even when they are left uncovered, the interest parity condition holds rather well. Given that large deviations from the UIP-condition have commonly been found with modern data—which has created the so-called UIP-puzzle—this is a remarkable result. Between the chief financial centres of Europe, no such puzzle seems to have existed during the classical gold standard.

The paper is organised as follows. The next section provides the historical background in terms of discussing the behaviour of the exchange rate during the gold standard, how short and long-bills of exchange determined the relevant exchange and discount rates, as well as the status of sterling as international currency and London as global financial centre. Section 3 provides the theoretical background comparing the specification of the covered and uncovered interest parity condition during the classical gold standard with the standard (or textbook) version. Section 4 discusses the discount and exchange rate data for the major European currencies and the decades before the First World War. Section 5 introduces the econometric strategy and presents the results. Section 6 summarises and concludes.

## 2 Historical background

As mentioned at the outset, interdependencies between exchange and interest rates are not an exclusive phenomenon of today's financial markets. As early as the $19^{th}$ century, bankers (see e.g. Clare, 1895) and theorists (see e.g. Goschen, 1861) were apparently aware of the economic forces tying the short-term return of different currencies together. Perhaps, the reason why this seems to have been forgotten is that the currency system, the instruments used in international finance, and the structure of the international financial system differed sufficiently to hide the at the time obvious relationships to the modern eye. Therefore, the following paragraphs briefly present some of the pillars on which the connection between exchange and interest rates during the classical gold standard were based.

## 2.1 Exchange rates during the gold standard

From around 1880 until the outbreak of the First World War in 1914, the gold standard served as role model for the international currency system. Though the definition of the value of a currency in terms of gold—the so-called mint-par—gave rise to officially fixed parities, the rates on the foreign exchange markets fluctuated nevertheless noticeably. A reason was that international gold shipments, which enforced the mint-par, were costly. Hence, the (sight-bill) exchange rate could fluctuate within a band whose width was roughly delimited by the so-called gold-points. The left panels of Figure 3 of Section 4 illustrate this for several currencies and reveal that, compared with nowadays, exchange rates were much more stable. It was primarily the prospect of international gold transactions that forestalled deviations of more than 1 per cent from the mint-par. As long as gold was allowed to be moved more or less freely and central banks stood ready to convert their currency into gold, a self-correcting mechanism anchored the exchange rate at the mint-par.

An ongoing debate is still trying to determine the gold-points. There are several reason why this has turned out to be challenging including the numerous and time-varying cost components to transport gold across countries (Clare, 1902, pp. 78-79), the complications when triangular gold arbitrage is possible (Coleman, 2007), and the practice of some central banks to implement monetary policy by, among other things, slightly manipulating the official gold price (Ugolini, 2013). To nevertheless provide some rough values, for the French franc, contemporaneous sources set the upper and lower gold-points at, respectively, around 25.12 Fcs/£ and 25.32 Fcs/£ (Clare, 1895, p.126), for the Dutch guilder at around 12.05 Fl/£ and 12.15 Fl/£ (Tate, 1908, p.328), and for the German mark at around 20.32 M/£ and 20.53 M./£ (Clare, 1895, p.131). However, different values have appeared in the literature depending e.g. on whether the gold-points are directly calculated from transaction costs or indirectly from econometric techniques estimating a latent threshold. For a recent discussion on this, see Bignon *et al.* (2015).

Even in its heyday, the gold standard was far from being a homogenous currency system. Only a handful of countries—including Britain, France, Germany, and the Netherlands—came close to the theoretical ideal of a freely convertible monometallic currency backed by gold (see Eichengreen, 2008, pp.20ff.). Across these relatively wealthy and financially advanced European nations, anchoring the official exchange rate at the mint-par meant that the remaining exchange rate fluctuations were, arguably, almost entirely the result of international differences in interest (discount) rates, or what Clare (1902, p.94) calls the value of money. Conversely, other countries around the world restricted the convertibility of their currency into gold, retained elements of silver or bimetallic standards, or used even inconvertible paper-money. The exchange rates of these countries were less stable (Einzig, 1962, 198-199) and responded heavily to fluctuations in international trade or changes in foreign indebtedness (Clare, 1902, p.94).

## 2.2 Bills of exchange were a key financial instrument

Amid an era of widespread economic and political stability, the second half of the $19^{th}$ century witnessed an unprecedented expansion of cross-border trade and capital flows (Obstfeld and Taylor, 2004). However, owing to the costs as well as the inelastic supply of gold backed money, the necessary increase in international payments would probably not have happened, had most transactions been settled by means of precious metal. The required international capital flows exceeded by far the volume of trade, which was, in turn, far larger than cross-border transfers of gold (Eichengreen, 2008, pp.24ff.). This reflects that, from the Middle Ages onward, bills of exchange were commonly used to finance international payments (Denzel, 2010 ch.3, Einzig, 1962, ch.7). In essence, a bill of exchange was a written order by an issuer, called the drawer, instructing a counterparty[1], called the drawee, to pay a certain amount of money at a specific place either immediately (sight-bill) or at the end of a given—usually three months—term to maturity (long-bill).

Bills could be issued on a foreign country or city with a different currency. Since they dominated international payments during the $19^{th}$ century, bills of exchange determined the relevant foreign currency price for cross-border business (see e.g. Denzel, 2010, ch.3.3). As such, the market exchange rates quoted for foreign bills could deviate from the mint-par. More specifically, when the demand for bills on a foreign currency was relatively high and/or the supply relatively low, the market exchange rate of that currency appreciated. What will be important for interest parity considerations is that fluctuations of the market exchange rate occurred with sight as well as with long-bills.

By combining elements of credit with money transfers, bills of exchange became the preferred instrument to arrange cashless payments. Though originally designed to finance trade (trade-bill), around 1900, bills of exchange were also widely used for purely financial purposes

[1]Cheques are a specific type of a bill in terms of being a written order to a *bank* to pay a specified amount upon presentation (Cassis, 2010, p.296).

(finance-bill) (Goschen, 1861, ch.3; Escher, 1913, ch.2; Clare, 1895, ch.13). It were in particular banks that recognised that bills of exchange provided an excellent instrument for short-term foreign investments. As an alternative to the local money market, e.g. banks of the principal countries on the European continent reinvested substantial parts of the savings deposited with them in long-bills issued in London. By way of contrast, before the First World War, British banks by and large ignored foreign bills (Clare, 1895, p.89; Clare, 1902, p.95; Thomas, 1921, p.78).

Especially when bills of exchange were drawn on a bank with a good reputation, they were seen as safe investment or "first-class" asset. This provided the basis for turning a bill into a transferable financial instrument, which could be sold well before its due date to a third party, called acceptor (often a discount house), taking over the responsibility to make the final payment (Cassis, 2010, p.84). From the perspective of the drawer, the selling (or discounting) of bills had the advantage of receiving early payment, but came at the price of the so-called discount-rate, that is the interest charged by the acceptor. The development of discount markets and specialised discount houses meant that bills of exchange became tradable and, hence, an "admirably liquid security" (Spalding, 1915, p.80). Owing to the pivotal role of bills of exchange within the financial system, the discount rate became one of the most closely watched interest rates during the $19^{th}$ century.

## 2.3 London serves as global financial hub

Reflecting the role of Britain as leading industrial nation, during the four decades preceding the First World War, London had established itself as principal hub for arranging, funding, and insuring the bulk of international trade and payments (Cassis, 2010, pp.83ff.). Though other financial centres—in particular, Paris, Berlin, or Amsterdam in Europe and New York in America—also witnessed rapid developments, the dominance of London was such that, prior to 1914, around half of global trade was financed by bills denominated in sterling, which had obtained the status of international currency par excellence (Atkin, 2005, p.5). With bills of exchange accounting for most international payments, this implied that a group of British merchant banks and discount houses accepted and discounted vast amounts of sterling-bills (generally of three months maturity) sustaining the most liquid money market in the world. Banks of the European continent, through their London branches or agents, took part in this market (Thomas, 1921, pp.80ff.), and thanks to the relatively unhindered flow of capital and the technological progress in telecommunication (telegraph, telephone), financial centres became closer intertwined and a genuine international capital market began to emerge (Cassis, 2010, p.131). However, with the sterling-bill constituting an almost globally accepted security, London struggled to develop a sizable foreign exchange market. Hence, at the dawn of the $20^{th}$ century, it were rather the above-mentioned continental capitals where the largest foreign exchange markets were located (Atkin, 2005, ch.1). French or German firms were indeed in constant need to discount large amounts of sterling-bills they received from exporting goods or capital, and drew similarly large amounts of sterling-bills from local banks to fund imports.

Against this background, the interest and exchange rates set in London served as international landmark. The corresponding data were published, typically on a weekly basis, in the financial press in Britain and abroad. Figure 1 provides examples taken from the $3^{rd}$ of March 1888 edition of The Economist. The top panel shows the London Course of Exchange bulletin, which reports the exchange rates on various foreign cities for the two most recent trading days (here $28^{th}$ of February and $1^{st}$ of March 1888) at the Royal Exchange, which was the principal market for foreign bills in Britain (see Clare, 1895, ch.8). For each foreign city, two quotations are given. The first (better) rate refers to "first-class paper", which generally meant bills of exchange involving banks with a good reputation. The second (higher) rate applied to ordinary trade bills involving little-known firms. When contemplating the actual exchange rate data in Section 4, it will be important to remember that banks preferred first-class bills (Thomas, 1921, ch.8; Clare, 1895, p.90; Clare, 1902, p.98f.). The reason was that they were highly liquid and a widely accepted, safe asset. Of note, most exchange rates in London refer to three months bills. Bearing witness to the minor importance of sight-bills in the London market, sight or cheque-rates were often only quoted on Paris or Amsterdam (Clare, 1902, pp.82, 85). Conversely, according to the middle panel, sight-rates quoted abroad *on* London existed for several European financial centres including Paris, Amsterdam, and Berlin. Finally, the London and continental discount rates (Clare, 1910, pp.94ff; Spalding, 1915, pp.80ff.) were also published in The Economist (see bottom panel of Figure 1).

Figure 1: Exchange and interest rates on the 1$^{st}$ of March 1888

**London Course of Exchange**

(Bulletin published in The Economist, 3rd March 1888)

| On. | Usance. | Price Negotiated on 'Change. Feb. 28. | | Mar. 1. | |
|---|---|---|---|---|---|
| Paris | Cheques | 25 25 | 25 30 | 25 27½ | 25 32½ |
| Ditto | 3 months | 25 42½ | 25 47½ | 25 43¾ | 25 48¾ |
| Amsterdam | ” | 12 2¾ | 12 3¼ | 12 2¾ | 12 3¼ |
| Ditto | At sight | 12 0¾ | 12 1¾ | 12 0¼ | 12 1¼ |
| Berlin | 3 months | 20 49 | 20 53 | 20 48 | 20 52 |
| Hamburg | ” | 20 48 | 20 52 | 20 48 | 20 52 |
| Frankfort | ” | 20 48 | 20 52 | 20 48 | 20 52 |
| Vienna and Trieste | ” | 12 85 | 12 90 | 12 82½ | 12 85 |
| Antwerp | ” | 25 45 | 25 50 | 25 45 | 25 50 |
| Petersburg | ” | 19 | 19½ | 19 | 19½ |
| Genoa, Naples, &c. | ” | 26 07½ | 26 12½ | 26 07½ | 26 12½ |
| Madrid, Barcelona, &c. | ” | 46⅛ | 46¼ | 46⅛ | 46¼ |
| Lisbon and Oporto | ” | 52⅝ | 52¾ | 52⅝ | 52¾ |

**Foreign rates of exchange on London**

(Bulletin published in The Economist, 3rd March 1888)

| | Latest Dates. | Rates of Exchange. | Usance. | | Latest Dates. | Rates of Exchange. | Usance. |
|---|---|---|---|---|---|---|---|
| Paris | Mar. 1 | 25 30 | Cheques | Rio Janeiro | Mar. 1 | 24 5/16 d | 90 dys st |
| Amsterdam | — 1 | 12·05 | Short | Buenos Ayres | Feb. 4 | 48¾d | ” |
| Berlin | — 1 | 20·36½ | ” | Do | Mar. 1 | 48½% | |
| Do | — 1 | 20·27½ | 3 mos. | Montevideo | Feb. 4 | 51¾d | ” |
| Hamburg | — 1 | 20·27 | ” | Valparaiso | — ... | ... | ” |
| Frankfort | — 1 | 20·38 | Short. | Melbourne | Jan. | ⅜ ⅝ p.sg. | 60 dys st |
| Vienna | — 1 | 12·08 | ” | Sydney | — | ⅜ ⅝ ds.bg | ” |
| St. Petersburg | — ... | ... | 3 mos. | Adelaide | — | | ” |
| Constantinople | — 1 | 111·75 | ” | Calcutta | Mar. 2 | 1/4 11/16 | telegra'c |
| Genoa, &c. | — ... | | ” | Bombay | — 2 | 1/4 11/16 | transf'rs |
| Madrid | — 1 | 25·58 | ” | Hong Kong | — 2 | 3/1½ | 4 m. sgt. |
| New York | — 1 | 4·85¼ | 60 dys st | Shanghai | — 2 | 4 3⅝ | ” |

**Discount quotations current in the chief continental cities**

(Bulletin published in The Economist, 3rd March 1888)

| | Bank Rate. % | Open Market. % | | Bank Rate % | Open Market % |
|---|---|---|---|---|---|
| Paris | 2½ Feb. 16, '88] | 2¼ | Vienna | 4 Jan., '88] | 3⅛ |
| Berlin | 3 May 15, '87] | 1½ | Genoa | 5½ Dec. '86] | 5 |
| Hamburg | 3 | 1⅜ | St Petersbrg. | 5 Dec. 12, '85] | 6½ |
| Frankfort | 3 | 1⅝ | Madrid | 4 | 4 |
| Amsterdam | 2½ May 29, '85] | 2¼ | Lisbon | 5 Apl 86] | 4½ |
| Brussels | 2½ Feb. 23, '88] | 2⅜ | Copenhagen | 3 | 3 |

# 3 Interest parity conditions now and then

Formal discussions about the interest parity condition can be traced back to the comparisons of the yield on various forms of Indian debt in Irving Fishers' *Appreciation of Interest* (1896, ch.9). In particular, Indian bonds were partly denominated in gold and partly in silver currency. This observation gave rise to the hypothesis that interest rate differences between these bonds should, on average, be offset by exchange rate changes between the corresponding currency standards. Along these lines, but for monometallic currencies based on gold, the top panel of Figure 2 compares the sterling (£) return in London with the return on French francs (Fcs.) in the Paris money market. In particular, departing from the top left box and following the black arrows, an arbitrageur (typically a bank) could either invest in the London money market at the three-months discount rate $i_t$, or abroad requiring first a conversion of sterling into French francs at the London on Paris sight-rate $S_t$, then an investment in the Paris money market yielding the discount rate $i_t^*$, and upon maturity (at date $t+m$) a conversion back at the expected sight-rate of Paris on London $E[S_{t+m}^*]$. Reflecting a situation where the yield of $1 + i_t$ in London and of $(1 + i_t^*)(S_t/E[S_{t+m}^*])$ on the Continent—including an expected currency gain/loss of $S_t/E[S_{t+m}^*]$—are equilibrated, the standard UIP-condition is given by

$$(1 + i_t) = (1 + i_t^*)\frac{S_t}{E[S_{t+m}^*]}. \tag{1}$$

A corresponding condition can also be derived from the perspective of Paris with French francs being the local currency. Adopting the conventional quotation at the time[2] and denoting the sight-bill rate in Paris on London by $S_t^*$, the grey arrows of the top panel of Figure 2 lead to the continental version of the UIP-condition, that is

$$(1 + i_t^*) = (1 + i_t)\frac{E[S_{t+m}]}{S_t^*}. \tag{2}$$

Of course, (1) and (2) do not differ if the same exchange rate is quoted across financial centres, meaning $S_t = S_t^*$ and, by extension, $E[S_{t+m}] = E[S_{t+m}^*]$. In modern work on the interest parity condition, this assumption is usually made to reflect that in today's globalised financial system, even tiny differences in spot exchange rates are quickly traded away. By way of contrast, around 1900, the number of financial firms engaging in international arbitrage

[2] The London and continental exchange rates were both expressed in terms of sterling per foreign currency (see Clare 1898 ch.9).

was arguably still relatively small (Cassis, 2010, pp. 131ff.). This is perhaps the reason why different sight-rates for bills *in* and *on* London were listed in the course of exchange bulletins (compare Figure 1). Moreover, across the various financial centres, the exchange rate quotations were highly incomplete (see Flandreau and Jobst, 2005) and depended also on the quality of a bill. Taken together, these observations suggest that, during the classical gold standard, there is no guarantee that the sight-rates quoted in different financial centres were typically equivalent.

Since the future exchange rate is a priori unknown, UIP-transactions are inherently risky. It is well-known that the covered, or non-speculative, version of interest parity relies on contractual agreements on future currency conversions to eliminate all exchange rate risk. Amid an expansion of foreign exchange trading on organised forward markets, Keynes' *Tract on Monetary Reform* (1923, pp.115-149) popularised the idea of a covered interest parity (CIP) condition, where the expected exchange rate $E[S_{t+m}]$ is replaced with the forward rate $F_{t+m|t}$ (see Isard, 1995, ch.5). Hence, (1) and (2) become

$$(1+i_t) = (1+i_t^*)\frac{S_t}{F_{t+m|t}^*} \tag{3}$$

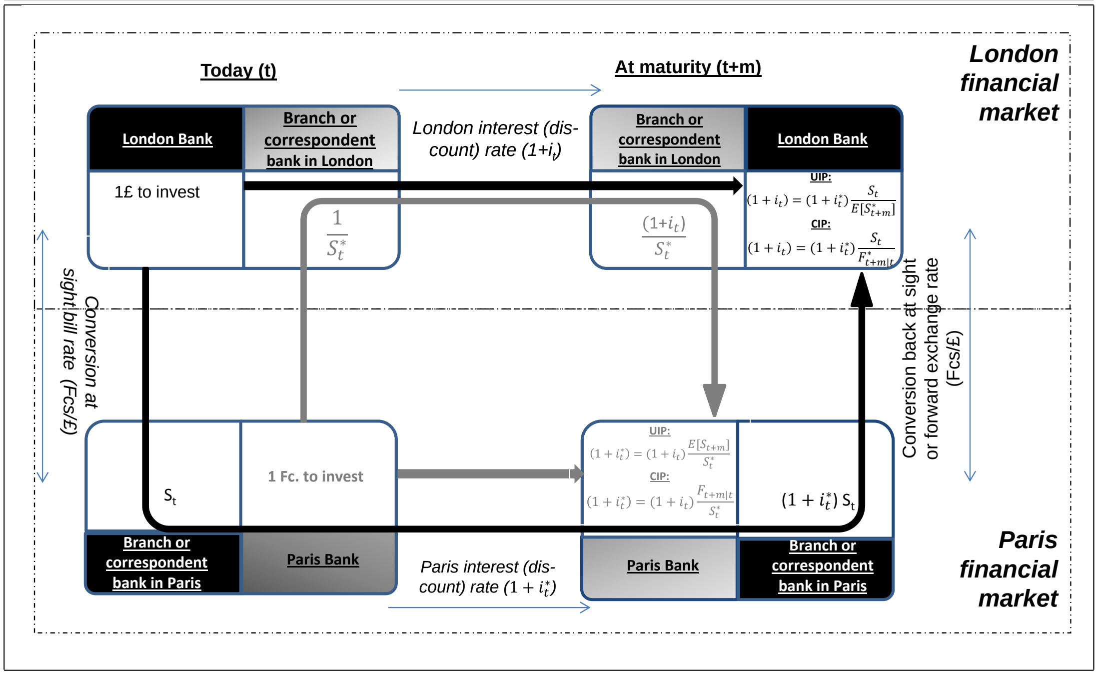
$$(1+i_t^*) = (1+i_t)\frac{F_{t+m|t}}{S_t^*}. \tag{4}$$

Equations (3) and (4) represent the standard form of covered interest arbitrage involving financial markets, on which currencies can be traded for future delivery against payment on delivery at a prearranged exchange rate. However, for most currencies—and namely for sterling—regular forward exchange markets did not develop until after the First World War (Einzig, 1962, pp.182-183; Einzig, 1967, pp.7-8).[3] Instead, early strategies to exploit international differences in interest rates relied on investments in long-bills (see e.g. Clare, 1895, ch.14-16; Spalding, 1915, pp.80ff.; Thomas, 1921, pp.78ff.), rather than forward exchange transactions, which represent a higher stage of financial development (Einzig, 1967, p.6).

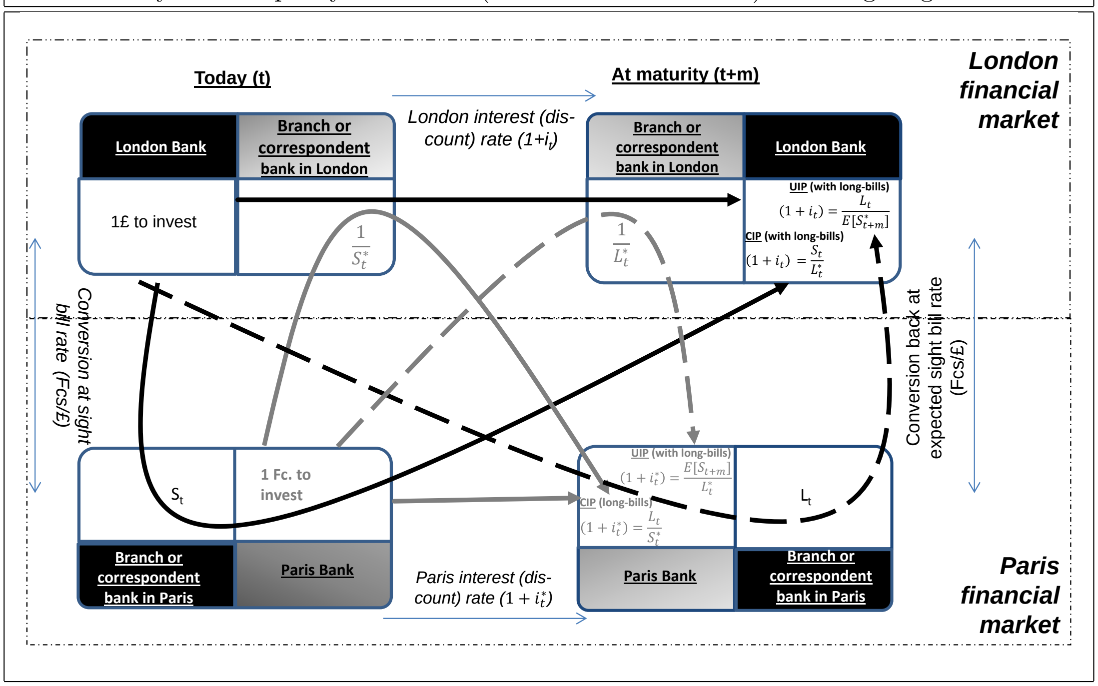
[3]During the second half of the $19^{th}$ century, an early forward market had already developed in Vienna to deal with the instable behaviour of the Austrian-Hungarian currency (see Einzig 1967, pp. 7-8), which was effectively on a flexible exchange rate until the 1890s (Flandreau and Komlos, 2006). It is therefore, perhaps, not surprising that for the case of Austria-Hungary, a verbal analysis of covered interest arbitrage transactions between Vienna and Berlin can already be found in Lotz (1889, pp.1277ff.). Flandreau and Komlos (2006) provide empirical results on the CIP transactions between Berlin and Vienna between 1896 and 1914. Furthermore, around 1900, another active forward exchange market existed for the Russian rouble (Einzig, 1962, p.214), which also consisted largely of inconvertible paper money with a relatively unstable exchange rate (Clare, 1895, p.159-160).

Figure 2: Standard and early (gold standard) interest parity conditions

Standard interest parity conditions (covered version involves forward exchange market)




Early interest parity conditions (uncovered and covered) involving long-bills




Before going into the details, it is important to realise that, though long-bills were used to hedge against exchange rate risk (Einzig, 1967, p.6), they differ from a modern forward contract in terms of not only embodying a future exchange rate, but also giving rise to a short-term credit (compare Section 2.2). Since bills of exchange transactions require an immediate payment (the price of the bill) from the buyer, who is promised to receive a certain amount of foreign currency back upon maturity (see Clare, 1898, ch.7; Clare, 1902, p.82), interest parity conditions involving long-bills differ slightly from their modern counterparts. In particular, turning first to the covered version, as illustrated by the solid black arrows in the bottom panel of Figure 2, e.g. a London based bank could first buy a sight-bill on a foreign city such as Paris at the sight-rate $S_t$ and, to safeguard against exchange rate risk, combine this with an immediate investment in a long-bill in Paris on London at the long-rate $L_t^*$. To eliminate arbitrage opportunities, the margin $S_t/L_t^*$ between these rates had, on average, to be equal to the return $1 + i_t$ in the London discount market, that is

$$\frac{S_t}{L_t^*} = (1 + i_t). \tag{5}$$

Then again, following the solid grey arrows of Figure 2, a similar condition can also be expressed from the Continent's perspective, which yields

$$\frac{L_t}{S_t^*} = (1 + i_t^*). \tag{6}$$

Furthermore, investments in sight and long-bills could also have been combined to give rise to UIP-transactions. For this, a London bank would have to first acquire a long-bill on, say, Paris. Upon maturity, the amount of foreign currency payable had to be invested in a sight-bill issued in Paris *on* London at the expected sight-rate $E[S_{t+m}^*]$. For an interest parity to emerge, the resulting return had to coincide with the return of $1 + i_t$ in the London discount market. As becomes clear by following the *dashed* black arrow in the bottom panel of Figure 2, from the perspective of London, the long-bill version of the UIP is given by

$$\frac{L_t}{E[S_{t+m}^*]} = (1 + i_t). \tag{7}$$

Finally, by following the dashed grey arrow, the same condition can again be expressed from the perspective of the Continent, which yields

$$\frac{E[S_{t+m}]}{L_t^*} = (1 + i_t^*). \tag{8}$$

Note that in (7) and (8), the interest parity conditions are uncovered, because the actual return of the long-bill investment depends on the, a priori unknown, rate of a sight-bill at future date $t + m$.

In sum, after solving (5) to (8) for $L_t$ or $L_t^*$, the long-bill versions of the CIP and the UIP concur nicely with the historical observation that the price of a bill payable at a future date was "based upon the sight-rate, rising and falling in agreement with it [...]" (Clare, 1902, p.83) as well as the principle that—owing to the immediate capital outlay associated with a long-bill—holders had to be compensated for the opportunity costs of awaiting payment in terms of an implicit interest rate (Goschen, 1861, pp.52ff.; Clare, 1895, ch.12; Clare, 1902, pp.82ff.; Spalding, 1915, ch.6).[4]

# 4 Weekly discount and exchange rate data

From tables such as those of Figure 1, for the decades before the First World War, Neal and Weidenmier (2003) have compiled weekly time series of interest (discount) and exchange rates across a range of financial centres. However, the exchange rate data are only available on a highly selective basis. As has already been said, most London rates referred to long-bills with a three-months term to maturity. Conversely, the chief financial markets on the European continent (Paris, Berlin, Vienna, or Amsterdam) typically quoted sight-rates on London, whilst elsewhere long-rates on London with a 60 days or three-months "usance" were common. Finally, owing to the lack of organised forward exchange markets, forward rates were not quoted. Taken together, this implies that only some of the interest parity conditions of the previous section can be confronted with the data. Recall from the discussion above that by far the most common way to earn foreign interest was to invest in long-bills linking the London money market with other financial centres across Europe. More specifically, for the 1880 to 1914 period (the common sample ends in December 1913), the data include money market discount rates in Paris, Amsterdam, and Berlin (denoted here by $i_t^*$) as well as in London (denoted here by $i_t$). Furthermore, for these continental financial centres, there are sight-bill rates on London ($S_t^*$) and (with the exception of Berlin) in London ($S_t$) as well as long-rates for three-months bills in London, that is $L_t$. Conversely, a long-rate on London, denoted by $L_t^*$, was only published in Berlin. Table 5 of the appendix provides an overview of the definition and sources of these variables.

Contemplating first the case of Paris, which was the most important financial centre after London (Cassis, 2010, pp.101ff.), the top left panel of Figure 3 depicts the sight-rates of

[4]The stamp duty levied on certain transactions impacted also upon the price of long-bills. Clare (1895, ch.12) shows how the London bill stamp, which amounted to less than 0.1 per cent of the bill's value, was priced into the long-rate (together with other minor expenses).

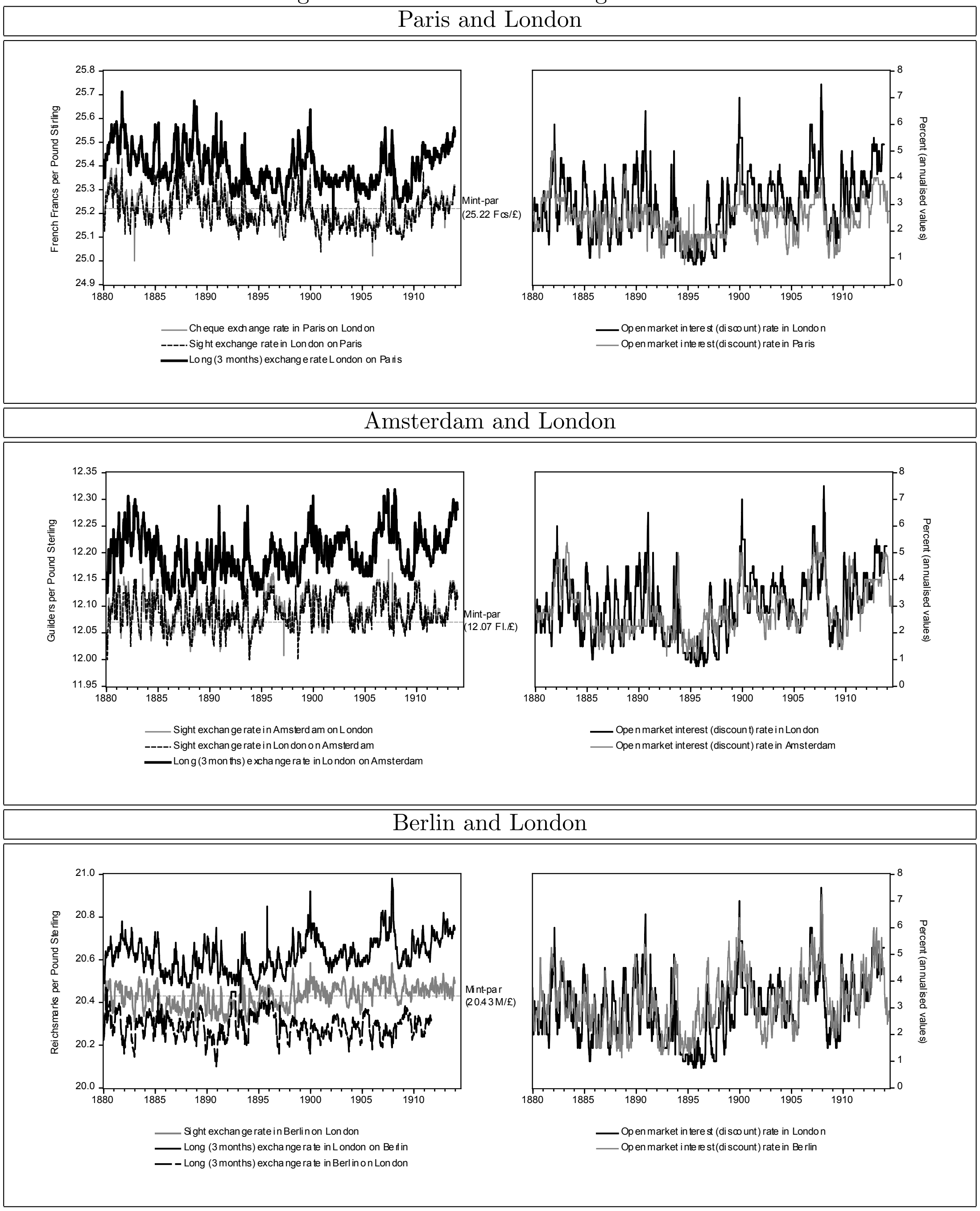
London on Paris and Paris on London as well as the long-rate of London on Paris. Since banks invested preferably in first-class paper, these rates refer to the lower value reported in Figure 1. As mentioned above, owing to the gold standard, the sight (or cheque) rates fluctuated around the mint-par of 25.22 Fcs./£. Though the rapid transmission of financial information by telegraph implied that sight-rates quoted in different financial followed each other quite closely (Cassis, 2010, p.76; Clare, 1902, p.99), the rate in Paris deviated still frequently from the corresponding London quotation.[5] Furthermore, the long-rate of exchange gives implicitly rise to an interest rate in the sense of being above the sight-rate. A buyer of a long-bill in London could, hence, almost be certain to receive more French francs than he would have to lay out for a sight-bill in three months time. Note that a long-rate in Paris on London is not available. The top right panel of Figure 3 depicts the interest rates from discounting a bill in the open money markets of London or Paris. Though these rates follow each other quite closely, the London rate was more volatile.

Between 1880 and 1914, The Economist provides another uninterrupted set of time series with exchange rates and the discount rate for the Dutch guilder, whose mint-par stood at 12.07 Fl./£ throughout those years. Though being well beyond its golden age, Amsterdam still punched above its weight as international financial centre (see Cassis, 2010, p.125; Einzig, 1962, p.177) and Holland is mentioned in Clare (1902, p.94) as one of the countries with a considerable investment demand for London bills. Accordingly, compared with the French franc, the middle panel of Figure 3 shows the familiar behaviour of the discount and the guilder-sterling exchange rates.

In contrast to the relative decline of Amsterdam, during the 1890s, Berlin became an increasingly more important financial centre (Cassis, 2010, pp.108ff.; Flandreau and Jobst, 2005, p.989). The bottom panel of Figure 3 reports the data for the German mark. The case of Berlin differs slightly in the sense that long-rates existed both in Berlin and in London, whilst only a sight-rate is published in Berlin on London (see also Figure 1). Note that, in compensation for providing a short-term credit, the Berlin long-rate lies below the corresponding sight-rate to make sure that the holder of a long-bill on London could almost be certain to receive more pounds than he would have to lay out for a sight-bill in three

[5]Indeed, for weekly data between 1880 and 1913, a regression of the Paris on London ($S_t^*$) onto the London on Paris $S_t$ sight-rate yields $S_t = 1.69 + 0.93S_t^*$ (Nr. of obs=1772; $R^2$=0.95). With a standard deviation (that is robust to heteroscedasticity and autocorrelation by applying the Newey-West correction) of 0.01, the slope coefficient differs from a value of 1 at any conventionally used level of rejection. This result lends support to the discussion of Section 3 that, during the gold standard, sight-rates quoted for the same currency in different financial centres could (even on average) deviate from each other.

Figure 3: Discount and exchange rates




months time.

For the other financial centres around the world, exchange rates in and on London were only published in a highly incomplete manner, or even not reported at all. This does not necessarily mean that these rates did not exist. Rather, the non-disclosure suggested, arguably, that for some countries certain foreign exchange transactions were "out of the ordinary course" and, hence, there was "no recognized quotation" (Clare, 1902, p.82). Therefore, for most currencies, the UIP and CIP-conditions of Section 3 cannot be estimated[6] and, for several reasons, active interest parity transactions involving the London money market were more or less restricted to the most developed countries on the European continent (Clare, 1902, pp.94ff). In particular, to shift large amounts of capital chasing higher foreign returns, a sufficiently advanced and internationally integrated banking system was required. Around 1900, perhaps, only a handful of countries had reached such a stage of financial development. Moreover, it is likely that only in the wealthiest countries a considerable amount of savings had accumulated to be at the disposal of the banking industry for interest parity transactions. Finally, it is also noteworthy that the countries mainly associated with the investment demand for foreign bills were at the heart of the gold standard. An entrenched mint-par was maybe necessary to eliminate prohibitively high levels of exchange rate risk.[7]

# 5 Estimating and testing the interest parity conditions

This section takes the various forms of the interest parity condition beyond the verbal discussion of Section 3, and conducts statistical tests by means of econometric methods that were unavailable around 1900. Recall from Section 4 that, even for the most developed European financial centres, exchange and interest data were in those days only partially available. Therefore, the standard version of the UIP-condition, given by (1) and (2), can only be estimated for the cases of Paris and Amsterdam, for which sight-rates in and on London were published, but not for the Berlin financial market, where only a sight-rate on London is available.[8] Moreover, the standard specification of the CIP, given by (3) and (4),

[6] Patchy data are available for Belgium and Italy. However, for the Belgian franc, the long-rate in London refers to Antwerp whilst, with an interruption between 1886 and 1902, only a sight-rate for Brussels can be found. Being a member of the Latin Monetary Union, via which the French mint-par of 25.22 Fcs./£ was adopted by several countries on the Continent, the exchange rate of the Belgian franc was relatively stable. Italy provided a different case. Though officially a member of the Latin Monetary Union, Italy only loosely pegged the lire to gold (Eichengreen, 2008, p.17). Aside from a short period in the 1880s, the mint-par of 25.22 lire/£ was never officially instituted, wherefore the exchange rate fluctuated markedly. The Italian time-series data are also incomplete in the sense that gaps exist for the sight-rate, which refers to Italy in general, whilst the long-rate in London refers to Genoa and Naples. The conditions in Austria-Hungary (or Vienna) have already been discussed in footnote 3. Owing to the incompleteness of the underlying data, the cases of Austria-Hungary, Belgium, and Italy are not pursued further.

[7] Given that New York began to emerge as an important financial centre (Goodhart, 1969; Cassis, 2010, pp.114ff.), it is perhaps surprising that the United States are mentioned as one of the most active participants in the London money market. Though the United States were an important trading partner for Britain, the capital flows were nowhere near those towards and from the principal European countries. The transaction costs of shifting financial capital across the Atlantic were still comparatively high. Hence, "the relative value of money, which is so largely responsible for the movements of the chief European exchanges, produces scarcely any perceptible effect on the American rate" (Clare, 1902, p.130).

[8] Though it could simply be assumed that $S_t = S_t^*$, as discussed in Sections 3 and 4, the anecdotal and statistical evidence would be at odds with this assumption.

does not lend itself to estimation since organised forward exchange markets barely existed before the First World War. Instead, early interest parity transactions used to combine investments in short and long-bills of exchange. Since the latter were mainly quoted in the London financial market, it is the continental CIP-condition of (6)—which combines a sight-bill transaction on London with a long-bill transaction in London—and the London UIP-condition of (7)—which combines a long-bill transaction in London with a future sight-bill transaction on London—that can be confronted with the data. For the CIP and the UIP-condition, the following paragraphs go through the details of the estimation method and present the empirical results.

## 5.1 Results for CIP-regressions (long-bills specification)

To conduct empirical tests, the early CIP-condition needs transforming into a regression equation. Following the modern literature and applying a logarithmic transformation (with the corresponding sight and long-bill rates being denoted by lowercase letters), the regression equation approximating (6)—the case for which data are available—is given by

$$l_t - s_t^* = \alpha + \beta(i_t^*) + \epsilon_t \qquad \text{Continental CIP-regression (long-bill specification).} \tag{9}$$

Here, $\alpha$ is an intercept reflecting such things as time-constant exchange rate risk (compare Chinn, 2006, p.9), $\beta$ is a slope-coefficient reflecting in how far exchange and interest rates move in tandem, and $\epsilon_t$ is a statistical error term. A scenario where there are no covered interest arbitrage opportunities left in the foreign exchange and money markets implies that $\beta = 1$. Whilst the estimation of the intercept and slope coefficient in (9) is relatively straightforward, some econometric issues arise as regards the corresponding standard deviations. In particular, a dataset combining three months long-bills with observations that have a weekly frequency gives inevitably rise to overlaps within the sample. It is well-known that this introduces moving-average terms to the residuals, which invalidates the estimates of the standard deviations, even when they are "robust" thanks to the conventional (Newey-West) method to correct for autocorrelation. A crude way to avoid this problem is to simply drop all overlapping observations. However, in the present case, this would remove more than 90 per cent of the sample. Chinn (2006, pp.9f.) has developed a more sophisticated correction for serial correlation due to such overlap, which can be dealt with by a heteroscedasticity and autocorrelation (HAC) robust variance-covariance matrix towards a fixed length of serial correlation of up to twice the overlap (here $\pm$ three-months or around 90 days/7 days per week $\approx$ 13 weeks).[9] Of note, regardless the chosen specification of the HAC-standard errors, they do not change the coefficient estimates.

For the different continental financial centres for which data are available, Table 1 summarises the CIP-results (involving long-bills).[10] In particular, column (1) reports the OLS-estimates of (9) for the Paris financial market. The coefficient of 1.01 for the slope ($\widehat{\beta}$), which does not differ from a value of one at any conventionally used level of rejection, concurs almost perfectly with the prior of the (covered) interest parity condition. This finding lends support to the above-mentioned claim that, around 1900, short-term foreign investments in London bills of exchange aligned the discount and exchange rates between countries with a relatively sophisticated banking system and being both on the gold standard as well as financially integrated through substantial capital flows. Reflecting similar conditions for the Amsterdam financial market, the corresponding results in column (2) lend again empirical support to the CIP-condition involving long-bills in the sense of giving rise to an estimated slope coefficient $\widehat{\beta}$ that does not differ from one at any conventionally used level of rejection. In column (3), the results for Berlin give rise to somewhat larger—and at the 5 per cent level of rejection significant—deviations of the coefficient estimate from the interest parity priors. Perhaps, this result reflects the fact that Berlin only began to emerge as important foreign exchange centre around the 1890s (compare Section 4).[11]

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[9] Several choices have to be made before estimating standard errors that are robust to heteroscedasticity and autocorrelation. Here, a Bartlett-kernel is used and there is no pre-whitening of the residuals. Of note, changing these options, or even altering the bandwidth to slightly different values than 13, did not overturn the essence of the results.

[10] The conventional unit-root tests (ADF, Phillips-Perron) suggest that even at the 1 per cent level of rejection, the time series of (9) are stationary (for the sake of brevity, the corresponding results are not reported here, but are available on request). Stationarity implies that no transformation of the data or testing for co-integration is warranted to estimate (9).

[11] The definition of the coefficient standard deviations is of crucial importance when trying to reject an interest parity condition. To get more precise estimates by accounting for possible contemporaneous correlations across currencies as well as currency-specific variances, Flood and Rose (1995) have advocated to pool the data and turn to seemingly unrelated regression estimation (SURE). Due to the data overlaps, the current observations from the gold standard era are also likely to exhibit correlation across time. Hence, as a robustness check, the results of Table 1 have been re-estimated with a period SURE. Similar to the case of individual time series, pooled unit root tests reject the hypothesis of an individual (see Maddala and Wu, 1999) and a common unit root (see Levin *et al.*, 2002). All specifications include cross-section specific effects. However, relaxing the restriction of having no correlation across the error component of the three financial centres (Paris, Amsterdam, Berlin) does not change the essence of the results. For the sake of brevity, the SURE-results are not reported here, but are available on request.

Table 1: CIP-results (early form of interest arbitrage via long-bills)

| | Paris | Amsterdam | Berlin |
|---|---|---|---|
| | (1) | (2) | (3) |
| Intercept ($\widehat{\alpha}$) | 0.003 | 0.01 | 0.01 |
| | (0.0004) | (0.001) | (0.001) |
| $i_t^*$ ($\widehat{\beta}$) | 1.01 | 0.97 | 0.96 |
| | (0.02) | (0.02) | (0.02) |
| Reject ($\beta = 1$ ) | | | ** |
| $R^2$ | 0.81 | 0.84 | 0.87 |
| N | 1,772 | 1,759 | 1,771 |

Notes: This table reports estimates of equation (9) with dependent variable $l_t - s_t^*$ (expressed as annualised value). Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. N denotes the number of observations. The null hypothesis that the CIP (via long-bill transactions) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

## 5.2 Results for UIP-regressions (long-bills and standard specification)

In contrast to the broad support for the CIP-condition (see e.g. Isard, 1995, pp.78-80), a large empirical literature has found results that are clearly at odds with the UIP-condition.[12] To uncover whether this so-called UIP-puzzle appears also with investments in long-bills issued in London on the European continent around 1900—e.g. the case for which the relevant data are available—(7) needs transforming into a logarithmic regression equation, that is

$$l_t - s_{t+m}^* = \quad \alpha + \beta(i_t) + \epsilon_t \qquad \text{London UIP-regression (long-bill specification).} \tag{10}$$

Equation (10) differs from (9) in terms of encapsulating exchange rate risk since the future sight-rate at which the return of a long-bill investment will be converted back is, a priori, unknown. Following the empirical literature on the UIP-condition, (10) assumes that ex-

[12] There are some caveats to the view that the CIP, but not the UIP, holds empirically. Arguably, the CIP-condition tends to break down during political or economic crises (Taylor, 1989). Furthermore, the UIP tends to work better between countries with fixed exchange rates (Flood and Rose, 1996); Chaboud and Wright (2005) have suggested that the UIP-condition holds during the very short-term, meaning when looking at intraday data and the overnight maturity; Chinn (2006) has suggested that the evidence against the UIP-condition is weaker when considering yields on government bonds with a five or ten year term to maturity; Lothian and Wu (2011) could not always reject the UIP-condition within a sample covering the very long-haul with two centuries worth of interest and exchange rate data; Herger (2016) could not reject the UIP-condition when employing panel data techniques and specifying the time-specific unobserved component as fixed effect.

pectations correspond with future exchange rates, or $E[s^*_{t+m}] = s^*_{t+m}$ (see e.g. Coleman, 2012, p.27). However, the null hypothesis that the UIP-condition holds requires that $\beta = 1$.

Employing the same econometric methodology as in Section 5.1, across the main financial centres on the European continent, Table 2 summarises the UIP-results involving a combined investment in long-bills issued in London with a sight-bill transaction on London upon maturity.[13] In particular, columns (1) to (3) report the OLS-estimates of equation (10) with Newey-West standard errors for comparisons of the London money market discount rate with bills of exchange transactions via Paris, Amsterdam, and Berlin. With estimates between 0.62 and 0.81, the slope coefficients are much closer to the hypothesised value than commonly found with modern data and slightly larger than the value of around 0.5 that Coleman (2012) has found when uncovered interest transactions were undertaken with the dollar-sterling exchange rate. Still, the hypothesis that $\beta = 1$ is clearly rejected.[14]

Table 2: UIP-results (early form of interest parity via long-bills)

| | Paris | A.dam | Berlin | Paris | A.dam | Berlin | Paris | A.dam | Berlin |
|---|---|---|---|---|---|---|---|---|---|
| Sample | ALL | | | Gold shipments | | | No gold shipments | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Intercept ($\widehat{\alpha}$) | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.02 | 0.02 | 0.01 |
| | (0.002) | (0.002) | (0.002) | (0.002) | (0.003) | (0.002) | (0.004) | (0.001) | (0.004) |
| $i_t$ ($\widehat{\beta}$) | 0.62 | 0.75 | 0.81 | 0.62 | 0.76 | 0.81 | 0.40 | 0.72 | 0.77 |
| | (0.05) | (0.07) | (0.06) | (0.05) | (0.08) | (0.07) | (0.12) | (0.08) | (0.12) |
| Reject ($\beta = 1$ ) | *** | *** | *** | *** | *** | *** | *** | *** | * |
| $R^2$ | 0.30 | 0.33 | 0.39 | 0.31 | 0.37 | 0.43 | 0.15 | 0.29 | 0.25 |
| N | 1,759 | 1,749 | 1,759 | 1,695 | 660 | 1,311 | 64 | 1,066 | 425 |

Notes: This table reports estimates of equation (10) with dependent variable $l_t - s^*_{t+m}$ (expressed as annualised value). The maturity $m$ equals three-months, that is 90 days or 90 days/7 days per week $\approx$ 13 weeks. Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. N denotes the number of observations. The null hypothesis that the UIP-condition (long-bill specification) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

Following Coleman (2012), the remaining columns of Table 2 split the sample according to whether or not imports or exports of gold bullion or coins with England took place during a given week.[15] As discussed in Section 2.1, around 1900, international gold transfers were typically seen as a mechanism correcting deviations from the mint-par. As regards the chief European financial centres, and in particular for Paris and Berlin, gold shipments from or towards London can be observed during most of the weeks between 1880 and 1914.

[13] The conventional unit-root tests (ADF, Phillips-Perron) suggest again that even at the 1 per cent level of rejection, the time series of all the regressions of this section are stationary (for the sake of brevity, the corresponding results are not reported here, but are available on request). The pooled unit root tests also reject the hypothesis of a common unit root.

[14] Similar to the discussion in the previous subsection, using a period SURE specification did not overturn the essence of the results.

[15] Import and export data of gold are published in the daily accounts of the Bank of England. A digital version of these accounts can be found at www.bankofengland.co.uk/archive/Pages/digitalcontent/archivedocs/dailyaccountbooks/dailyaccountbooks.aspx.

Therefore, the corresponding coefficient estimates in columns (4) to (6) deviate barely from the full sample in columns (1) to (3). Conversely, when no gold shipments took place, the estimated slope coefficients in columns (7) to (9) are slightly smaller, though not statistically different from those of the full sample.[16]

During the classical gold standard, large deviations of the sight-rate from the mint-par—and, in turn, large gold shipments—were usually associated with events of aggravated political or economic uncertainty (Clare, 1902, p.98). From a different angle, Table 3 addresses the question as to whether exchange and interest rates could react differently when the continental sight-rate $s_t^*$ of equation (10) deviates substantially from the mint-par, by distinguishing between observations within and outside (or on) the gold-points. Of note, the vagaries to determine the gold-points (compare Section 2.1) introduces an important caveat this approach. Still, based on the—admittedly rather arbitrary—values reported in Section 2.1, the slope coefficients are slightly larger for observations outside the gold points in Paris and Berlin (though the latter case encompasses only 43 observations). For the case of Paris, the value is even sufficiently close to one such that the UIP-condition cannot be rejected.[17]

Table 3: Gold points: UIP-results (early form of interest arbitrage via long-bills)

| | Paris | Amsterdam | Berlin | Paris | Amsterdam | Berlin |
|---|---|---|---|---|---|---|
| Sample: | Observations inside gold points | | | Observations outside (or on) gold points | | |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Intercept ($\widehat{\alpha}$) | 0.01 | 0.01 | 0.02 | 0.01 | 0.01 | -0.002 |
| | (0.002) | (0.002) | (0.002) | (0.01) | (0.005) | (0.01) |
| $i_t$ ($\widehat{\beta}$) | 0.58 | 0.73 | 0.77 | 0.81 | 0.69 | 1.34 |
| | (0.05) | (0.07) | (0.06) | (0.17) | (0.16) | (0.11) |
| Reject ($\beta = 1$ ) | *** | *** | *** | | * | *** |
| $R^2$ | 0.32 | 0.33 | 0.36 | 0.24 | 0.23 | 0.80 |
| N | 1,510 | 1,559 | 1,715 | 249 | 177 | 43 |

Notes: This table reports estimates of equation (10) with dependent variable $l_t - s^*_{t+m}$ (expressed as annualised value). The maturity $m$ equals three-months, that is 90 days or 90 days/7 days per week $\approx$ 13 weeks. Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. According to Section 2.1, the gold points of $s_t^*$ are 25.12 Fcs./£ and 25.32 Fcs./£ for France, 12.05 Fl./£ and 12.15 Fl./£ for the Netherlands, as well as 20.32 M/£ and 20.53 M/£ for Germany. N denotes the number of observations. The null hypothesis that the UIP (via long-bill transactions) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

In contrast to the covered continental investment in long-bills of the previous subsection, when leaving exchange rate risk uncovered, the results of Tables 2 and 3 do not suggest that around 1900 the realised return of $l_t - s^*_{t+m}$ of a London investment in a continental

[16] In his results, Coleman (2012) inserts a dummy variable to indicate whether gold exports occurred during a given week. However, inserting such a dummy variable designating observations with gold exports in Table 2 barely changed the coefficient estimates.

[17] For the sake of completeness, the results of the CIP regression of Table 1 have also been calculated distinguishing whether gold shipments took place and between observations within and outside the gold points. For the case of the CIP condition of (9), these distinctions did not matter.

financial centre coincided exactly with the discount rate $i_t$ in the London money market. This finding seems to concur with the observation, discussed in numerous guide books, that before the First Wold War, British banks were reluctant to invest in foreign bills (compare Section 2.2). Unfortunately, a clear reason why this was the case is not given and it is e.g. merely observed that ”foreign bills as an investment are strangely neglected by English bankers” (Clare, 1902, p.95) or that ”the continental banker is more eager to make a chance profit than the London banker, who is usually content with a safe and steady rate of interest” (Thomas, 1921, p.78). Perhaps, the key difference was that London investments in long-bills according to (10) were indeed subject to exchange rate risk, whilst this was not necessarily the case for the continental demand for London bills of exchange according to (9).

For transactions between London and Paris as well as Amsterdam—the financial centres for which sight-rates in and on London were regularly published during the 1880 to 1914 period—the nexus between discount and exchange rates can also be estimated via the standard specifications of (1) and (2) of the UIP-condition. With the usual logarithmic approximation, the corresponding regression equations are given by

$$s_t - s^*_{t+m} = \alpha + \beta(i_t - i^*_t) + \epsilon_t \quad \text{London UIP-regression (standard specification)} \tag{11}$$

$$s_{t+m} - s^*_t = \alpha + \beta(i^*_t - i_t) + \epsilon_t \quad \text{Continental UIP-regression (standard specification).} \tag{12}$$

In this form, the UIP-condition suggests that a relatively low (high) discount rate for a given currency is, on average, offset by a proportional appreciation (depreciation) of that currency. The regression equations (11) and (12) express this condition from the perspective of London and a continental city, respectively. Statistically, whether the UIP-condition holds is again a test on the the null-hypothesis that $\beta = 1$. Relying on the estimation techniques employed above, Table 4 suggest that the results based on the standard specification of the UIP-condition coincide, by and large, with those reported above. For foreign exchange transactions connecting the London financial market with Paris and Amsterdam, the slope coefficients of the UIP-regression are between 0.3 and 0.4 and differ significantly from the hypothesised value of one. However, slightly larger estimates arise with a sample containing only observations on or outside the gold-points (defined, as above, with respect to $s_t^*$). However, compared with the UIP involving long bills, the $R^2$ in Table 4 are relatively low suggesting that international differences in discount rates explained only a fraction of the changes of the sight-rate.

Taken together, in how far the UIP-condition can be rejected for transactions between the major continental financial centres and London around 1900 seems to depend on such things as the deviation from the mint-par. In any case, it should not be overlooked that a UIP-puzzle, in terms of massive deviations from the theoretically expected slope coefficient or even negative values for $\beta$ that have frequently been found with modern data (see e.g. Engel, 2014, p.495), does not arise in Tables 2 to 4. Concurring with the finding of Flood and Rose (1996), compared with the modern era with floating exchange rates between the major currencies, the UIP-condition held relatively well for the core countries linked through the fixed exchange rates associated with the classical gold standard between 1880 and 1914.

# 6 Summary and conclusion

By examining the interest parity condition as regards the investment demand for bills of exchange during the "belle epoque" of the gold standard (1880 - 1914), this paper is embedded in a vast literature about the financial determinants of exchange rates. The main innovation is to address this question with exchange rates between sterling and three European currencies (French franc, Dutch guilder, German mark) from the end of the $19^{th}$ and the beginning of the $20^{th}$ century. Two major observations can be taken away.

Firstly, as early as at the end of the $19^{th}$ century, the premise that interest and exchange rates are closely intertwined was prominently discussed in a number of guide books about the foreign exchanges and money markets. Though these books had a practical rather than a theoretical flavour, they nevertheless suggest that the idea of an interest parity condition is older than commonly thought. Above all, around 1900, it seems to have been widely-known that European banks exploit international differences in the (exchange rate adjusted) short-term return between the major gold-backed currencies. Yet, since international transactions occurred via different financial instruments (bills of exchange), and the financial system was structured around the gold standard, the links between the exchange and interest rates in those days are, maybe, no longer obvious to the modern eye.

Table 4: UIP-results (Standard specification of the UIP-regression)

| Sample: | All | | | | Inside gold points | | | | Outside gold points | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | London on | | Paris | A.dam | London on | | Paris | A.dam | London on | | Paris | A.dam |
| | Paris | A.dam | on London | | Paris | A.dam | on London | | Paris | A.dam | on London | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| Intercept ($\widehat{\alpha}$) | -0.004 | -0.002 | 0.0004 | -0.001 | -0.004 | -0.001 | 0.001 | -0.001 | -0.0001 | -0.01 | -0.004 | 0.003 |
| | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.0003) | (0.001) | (0.002) | (0.002) | (0.003) | (0.002) |
| $(i_t - i_t^*)$ ($\widehat{\beta}$) | 0.39 | 0.37 | | | 0.37 | 0.29 | | | 0.48 | 0.58 | | |
| | (0.07) | (0.07) | | | (0.07) | (0.07) | | | (0.19) | (0.20) | | |
| $(i_t^* - i_t)$ ($\widehat{\beta}$) | | | 0.37 | 0.33 | | | 0.34 | 0.25 | | | 0.47 | 0.66 |
| | | | (0.07) | (0.07) | | | (0.06) | (0.07) | | | (0.21) | (0.22) |
| Reject ($\beta$=1 ) | *** | *** | *** | *** | *** | *** | *** | *** | *** | ** | ** | * |
| $R^2$ | 0.10 | 0.07 | 0.09 | 0.06 | 0.10 | 0.05 | 0.09 | 0.04 | 0.08 | 0.14 | 0.07 | 0.14 |
| N | 1,759 | 1,748 | 1,760 | 1,747 | 1,510 | 1,559 | 1,511 | 1,569 | 249 | 177 | 249 | 178 |

Notes: This table reports estimates of equation (11) with dependent variable $s_t - s_{t+m}^*$ (expressed as annualised value) in columns (1), (2), (5), (6), (9), and (10) as well as of equation (12) with dependent variable $s_{t+m} - s_t^*$ (expressed as annualised value) in columns (3),(4), (7), (8), (11), and (12). The maturity $m$ equals three-months, that is 90 days or 90 days/7 days per week $\approx$ 13 weeks. According to Section 2.1, the gold points of $s_t^*$ are 25.12 Fcs./£ and 25.32 Fcs./£ for France, 12.05 Fl./£ and 12.15 Fl./£ for the Netherlands, as well as 20.32 M/£ and 20.53 M/£ for Germany. Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. N denotes the number of observations. The null hypothesis that the UIP-condition (standard specification) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

Secondly, for the most advanced European financial centres around 1900, regressing the short-term interest from discounting bills in the local money markets onto the return implied in foreign bills of exchange yields coefficients that are consistent with the covered, and sometimes even the uncovered, interest parity condition. This lends support to the historical claim that, at the time, the so-called investment demand for bills of exchange aligned the exchange rates of the major gold-backed currencies with the money market interest (or discount) rates. In view of the "puzzle" that the UIP-condition finds much less empirical support when looking at transactions between the major currencies with modern data, this is a remarkable result.

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# A Data appendix

Table 5: Description of the data set
Variables are collected for transactions between London and 3 continental financial centres (Paris, Amsterdam, Berlin) between 1880 and 1914. The data have a weekly frequency.

| Variable | Unit | Description | Source |
|---|---|---|---|
| $s_t$ | Continental currency per £ (in logarithms). | Exchange rate inherent in a sight-bill of exchange issued in London on a continental financial centre. The data are available for sight-rates on Paris and Amsterdam (but not for Berlin). For the case of Paris, the cheque rate has been used for $s_t$. | Neal-Weidenmier Gold Standard Database, Neal and Weidenmier (2003). Original sources are The Economist and the Commercial and Financial Chronicle. |
| $s_t^*$ | Continental currency per £ (in logarithms). | Exchange rate inherent in a sight-bill of exchange issued in a continental financial centre on London. The data are available for sight-rates in Paris, Amsterdam, and Berlin. For the case of Paris, the cheque rate has been used for $s_t^*$. | Neal-Weidenmier Gold Standard Database, Neal and Weidenmier (2003). Original sources are The Economist and the Commercial and Financial Chronicle. |
| $l_t$ | Continental currency per £ (in logarithms). | Exchange rate inherent in a long-bill of exchange issued in London on a continental financial centre. The term to maturity is three months. The data are available for long-rates on Paris, Amsterdam, and Berlin. | Neal-Weidenmier Gold Standard Database, Neal and Weidenmier (2003). Original sources are The Economist and the Commercial and Financial Chronicle. |
| $l_t^*$ | Continental currency per £ (in logarithms). | Exchange rate inherent in a long-bill of exchange issued in a continental financial centre on London. The term to maturity is three months. The data are available for long-rates in Berlin (but not in Paris and Amsterdam). | Neal-Weidenmier Gold Standard Database, Neal and Weidenmier (2003). Original sources are The Economist and the Commercial and Financial Chronicle. |
| $i_t$ | Per cent (annualised). | Interest on a short-term investment in the London money market. The interest arises from discounting a bill of exchange at the money market rate. | Neal-Weidenmier Gold Standard Database, Neal and Weidenmier (2003). Original sources are The Economist and the Commercial and Financial Chronicle. |
| $i_t^*$ | Per cent (annualised). | Interest on a short-term investment in the continental money market. The interest arises from discounting a bill of exchange at the money market rate. | Neal-Weidenmier Gold Standard Database, Neal and Weidenmier (2003). Original sources are The Economist and the Commercial and Financial Chronicle. |

# B Additional results (Not for publication)

## B.1 Unit root tests

Table 6: Unit root tests (pertaining to CIP long-bill version of Table 1)

**Unit Root Tests (Single variables)**

| | Paris | | Amsterdam | | Berlin | |
|---|---|---|---|---|---|---|
| | $i_t^*$ | $l_t - s_t^*$ | $i_t^*$ | $l_t - s_t^*$ | $i_t^*$ | $l_t - s_t^*$ |
| ADF | -5.97*** | -7.17*** | -4.53*** | -4.85*** | -5.26*** | -6.67*** |
| PP | -6.37*** | -14.8*** | -4.46*** | -11.0*** | -6.08*** | -8.71*** |

Notes: The null hypothesis is that there is a unit root. ADF is the Augmented-Dickey-Fuller test statistic. PP is the Phillips-Perron test statistic. All models contain an intercept. The lag-length has been selected via the Schwarz information criterion (SIC).

**Pool Unit Root Tests**

| Test Statistic | $i_t^*$ | $l_t - s_t^*$ |
|---|---|---|
| Common Unit Root | | |
| Levin, Lin, and Chu (2002) | -2.46*** | -5,89*** |
| Individual Unit Root | | |
| ADF - Maddala and Wu (1999) | 72.4*** | 104.6*** |
| PP - Maddala and Wu (1999) | 84.9*** | 313.2*** |

Notes: The null hypothesis is that there is a unit root. All tests statistics include an intercept. The lag-length has been automatically selected according to the SIC (starting from 20 lags). Details on the construction of the test statistics are given by the references.

Table 7: Unit root tests (pertaining to UIP long-bill version of Tables 2 and 3)

**Unit Root Tests (Single variables)**

| | Paris | | Amsterdam | | Berlin | |
|---|---|---|---|---|---|---|
| | $i_t$ | $l_t - s_{t+m}^*$ | $i_t$ | $l_t - s_{t+m}^*$ | $i_t$ | $l_t - s_{t+m}^*$ |
| ADF | -6.51*** | -9.88*** | -6.51*** | -5.40*** | -6.51*** | -8.56*** |
| PP | -5.91*** | -9.80*** | -5.91*** | -9.33*** | -5.91*** | -9.00*** |

Notes: The null hypothesis is that there is a unit root. ADF is the Augmented-Dickey-Fuller test statistic. PP is the Phillips-Perron test statistic. All models contain an intercept. The lag-length has been selected via the Schwarz information criterion (SIC).

**Pool Unit Root Tests**

| Test Statistic | $i_t$ | $l_t - s_{t+m}^*$ |
|---|---|---|
| Common Unit Root | | |
| Levin, Lin, and Chu (2002) | -0.85 | 0,92 |
| Individual Unit Root | | |
| ADF - Maddala and Wu (1999) | 37.4*** | 170.3*** |
| PP - Maddala and Wu (1999) | 30.7*** | 223.6*** |

Notes: The null hypothesis is that there is a unit root. All tests statistics include an intercept. The lag-length has been automatically selected according to the SIC (starting from 20 lags). Details on the construction of the test statistics are given by the references.

Table 8: Unit root tests (pertaining to UIP standard specification of Table 4)

| **Unit Root Tests (Single variables)** | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | London on Paris | | London on Amsterdam | | Paris on London | | Amsterdam on London | |
| | $i_t - i_t^*$ | $s_t - s_{t+m}^*$ | $i_t - i_t^*$ | $s_t - s_{t+m}^*$ | $i_t^* - i_t$ | $s_{t+m} - s_t^*$ | $i_t^* - i_t$ | $s_{t+m} - s_t^*$ |
| ADF | -9.24*** | -9.43*** | -8.31*** | -10.2*** | -9.34*** | -9.88*** | -8.31*** | -9.05*** |
| PP | -8.63*** | -10.7*** | -8.31*** | -12.1*** | -8.63*** | -10.7*** | -8.81*** | -11.9*** |

Notes: The null hypothesis is that there is a unit root. ADF is the Augmented-Dickey-Fuller test statistic. PP is the Phillips-Perron test statistic. All models contain an intercept. The lag-length has been selected via the Schwarz information criterion (SIC).

| **Pool Unit Root Tests** | | | | |
|---|---|---|---|---|
| Test Statistic | $i_t - i_t^*$ | $s_t - s_{t+m}^*$ | $i_t^* - i_t$ | $s_{t+m} - s_t^*$ |
| Common Unit Root | | | | |
| Levin, Lin, and Chu (2002) | -7.85*** | 15.0 | -7.85*** | 16.0 |
| Individual Unit Root | | | | |
| ADF - Maddala and Wu (1999) | 132.7*** | 161.3*** | 132.7*** | 151.4*** |
| PP - Maddala and Wu (1999) | 130.9*** | 206.1*** | 130.9*** | 204.7*** |

Notes: The null hypothesis is that there is a unit root. All tests statistics include an intercept. The lag-length has been automatically selected according to the SIC (starting from 20 lags). Details on the construction of the test statistics are given by the references.

## B.2 Additional results pertaining to Section 5.1

Table 9: SURE: CIP-results (early form of interest arbitrage via long-bills)

| | Paris | Amsterdam | Berlin |
|---|---|---|---|
| | (1) | (2) | (3) |
| Intercept ($\widehat{\alpha}$) | 0.003 | 0.01 | 0.01 |
| | (0.001) | (0.001) | (0.001) |
| $i_t^*$ ($\widehat{\beta}$) | 1.01 | 0.97 | 0.96 |
| | (0.06) | (0.02) | (0.01) |
| Reject ($\beta = 1$ ) | | | *** |
| $R^2$ | | 0.87 | |
| N | | 5,302 | |

Notes: This table reports estimates of equation (9) with dependent variable $l_t - s_t^*$ (expressed as annualised value). Estimation is by SURE with cross-section specific effects and allowing for cross-correlation in the residuals (across currencies and time). N denotes the number of observations. The null hypothesis that the CIP (via long-bill transactions) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

Table 10: Gold shipments: CIP-results (early form of interest arbitrage via long-bills)

| | Paris | Amsterdam | Berlin | Paris | Amsterdam | Berlin |
|---|---|---|---|---|---|---|
| Standard dev. | Gold shipments | | | No gold shipments | | |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Intercept ($\widehat{\alpha}$) | 0.003 | 0.01 | 0.01 | 0.003 | 0.01 | 0.01 |
| | (0.0004) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| $i_t^*$ ($\widehat{\beta}$) | 1.01 | 0.98 | 0.96 | 1.03 | 0.98 | 1.01 |
| | (0.02) | (0.02) | (0.02) | (0.06) | (0.02) | (0.04) |
| Reject ($\beta = 1$ ) | | | ** | | | |
| $R^2$ | 0.81 | 0.87 | 0.88 | 0.87 | 0.81 | 0.83 |
| N | 1,708 | 668 | 1,323 | 64 | 1,069 | 425 |

Notes: This table reports estimates of equation (9) with dependent variable $l_t - s_t^*$ (expressed as annualised value). Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. N denotes the number of observations. The null hypothesis that the CIP (via long-bill transactions) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

Table 11: Gold points: CIP-results (early form of interest arbitrage via long-bills)

| | Paris | Amsterdam | Berlin | Paris | Amsterdam | Berlin |
|---|---|---|---|---|---|---|
| Standard dev. | Observations inside gold points | | | Observations outside gold points | | |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Intercept ($\widehat{\alpha}$) | 0.003 | 0.01 | 0.01 | 0.004 | 0.01 | 0.02 |
| | (0.0003) | (0.001) | (0.001) | (0.001) | (0.002) | (0.005) |
| $i_t^*$ ($\widehat{\beta}$) | 1.02 | 0.98 | 0.97 | 0.97 | 0.96 | 0.72 |
| | (0.03) | (0.02) | (0.02) | (0.06) | (0.09) | (0.10) |
| Reject ($\beta = 1$ ) | | | * | | | *** |
| $R^2$ | 0.87 | 0.87 | 0.88 | 0.47 | 0.62 | 0.73 |
| N | 1,520 | 1,581 | 1,728 | 252 | 178 | 43 |

Notes: This table reports estimates of equation (9) with dependent variable $l_t - s_t^*$ (expressed as annualised value). Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. According to Section 2.1, the gold points of $s_t^*$ are 25.12 Fcs./£ and 25.32 Fcs./£ for France, 12.05 Fl./£ and 12.15 Fl./£ for the Netherlands, as well as 20.32 M/£ and 20.53 M/£ for Germany. N denotes the number of observations. The null hypothesis that the CIP (via long-bill transactions) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

## B.3 Additional results pertaining to Section 5.2

Table 12: SURE: UIP-results (early form of interest parity via long-bills)

| | Paris | Amsterdam | Berlin |
|---|---|---|---|
| | (1) | (2) | (3) |
| Intercept ($\widehat{\alpha}$) | 0.01 | 0.01 | 0.01 |
| | (0.0001) | (0.0001) | (0.0001) |
| $i_t$ ($\widehat{\beta}$) | 0.62 | 0.75 | 0.81 |
| | (0.002) | (0.003) | (0.002) |
| Reject ($\beta = 1$ ) | *** | *** | *** |
| $R^2$ | | 0.40 | |
| N | | 5,266 | |

Notes: This table reports estimates of equation (10) with dependent variable $l_t - s^*_{t+m}$ (expressed as annualised value). The maturity $m$ equals three-months, that is 90 days or 90 days/7 days per week $\approx$ 13 weeks. Estimation is by SURE with cross-section specific effects and allowing for cross-correlation in the residuals (across currencies and time). N denotes the number of observations. The null hypothesis that the UIP-condition (long-bill specification) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

Table 13: UIP-results (early form of interest parity via long-bills)

| | Paris | A.dam | Berlin |
|---|---|---|---|
| Sample | | Gold shipments | |
| | (1) | (2) | (3) |
| Intercept ($\widehat{\alpha}$) | 0.01 | 0.02 | 0.01 |
| | (0.002) | (0.003) | (0.002) |
| $i_t$ ($\widehat{\beta}$) | 0.61 | 0.72 | 0.79 |
| | (0.05) | (0.08) | (0.06) |
| Gold Export | -0.004 | -0.01 | -0.04 |
| | (0.001) | (0.001) | (0.001) |
| Reject ($\beta = 1$ ) | *** | *** | *** |
| $R^2$ | 0.32 | 0.39 | 0.45 |
| N | 1,668 | 655 | 1,307 |

Notes: This table reports estimates of equation (10) with dependent variable $l_t - s^*_{t+m}$ (expressed as annualised value). The maturity $m$ equals three-months, that is 90 days or 90 days/7 days per week $\approx$ 13 weeks. Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. N denotes the number of observations. The null hypothesis that the UIP-condition (long-bill specification) holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.

Table 14: Gold shipments. UIP-results (Standard specification of the UIP-regression)

| | Gold shipments | | | | No gold shipments | | | |
|---|---|---|---|---|---|---|---|---|
| | London on | | Paris | A.dam | London on | | Paris | A.dam |
| | Paris | A.dam | on London | | Paris | A.dam | on London | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Intercept ($\widehat{\alpha}$) | -0.004 | -0.001 | 0.0004 | -0.001 | -0.001 | -0.002 | -0.002 | -0.001 |
| | (0.0001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| $(i_t - i^*_t)$ ($\widehat{\beta}$) | 0.40 | 0.41 | | | 0.06 | 0.31 | | |
| | (0.07) | (0.10) | | | (0.16) | (0.07) | | |
| $(i^*_t - i_t)$ ($\widehat{\beta}$) | | | 0.37 | 0.36 | | | 0.08 | 0.29 |
| | | | (0.07) | (0.10) | | | (0.15) | (0.07) |
| Reject ($\beta$=1 ) | *** | *** | *** | *** | *** | *** | *** | *** |
| $R^2$ | 0.10 | 0.09 | 0.09 | 0.06 | 0.003 | 0.05 | 0.005 | 0.05 |
| N | 1,695 | 660 | 1,696 | 660 | 64 | 1,066 | 64 | 1,066 |

Notes: This table reports estimates of equation (11) with dependent variable $s_t - s^*_{t+m}$ (expressed as annualised value) in columns (1), (2), (5), (6), (9), and (10) as well as of equation (12) with dependent variable $s_{t+m} - s^*_t$ (expressed as annualised value) in columns (3),(4), (7), (8), (11), and (12). The maturity $m$ equals three-months, that is 90 days or 90 days/7 days per week $\approx$ 13 weeks. Estimation is by OLS with heteroscedasticity and autocorrelation robust (Newey-West) standard errors with a fixed bandwidth of 13 leads and lags. N denotes the number of observations. The null hypothesis that the UIP holds implies that $\beta = 1$. Significant deviations from this are indicated by * at the 10% level; ** at the 5% level, and *** at the 1% level.