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Preferred Habitat For Liquidity In International Short-term Interest Rates

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ABSTRACT

U.S. money market securities have been found to exhibit behavior consistent with preferred habitat for liquidity around year-ends (Griffiths and Winters (1997, 2004)). In particular, repurchase agreement and commercial paper yields tend to increase when the security begins to mature across the end of the year, and return to normal levels after the year-end obligations have been paid but before the calendar year-end. The competing hypothesis, window dressing by financial intermediaries around disclosure dates, requires that the increase in yields be sustained until after the turn of the year. This study is aimed at finding whether the behavior of international money markets around year-ends and quarter-ends is more consistent with preferred habitat for liquidity or window dressing. This is done by analyzing changes in LIBOR for different currencies around quarter-ends.

A second part of the study considers the effect of preferred habitat on the term structure of short-term interest rates. The expectations hypothesis of the term structure posits that future expected interest rates are implied by the current term structure. Empirical research suggests that the expectations hypothesis often does not hold, especially at the short end of the term structure. Preferred habitat for liquidity in short-term rates may be one of the reasons for the failure of expectations. The same LIBOR data set is used to test for the expectations in the presence of preferred habitat for liquidity. The empirical results of this study suggest that preferred habitat for liquidity in the short-term rates around quarter-ends and year-ends is not responsible for the failure of the expectations hypothesis in the data.

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LIST OF ABBREVIATIONS

ADF – Augmented Dickey-Fuller test

BBA – British Bankers' Association

CAPM – Capital Asset Pricing Model

CD – Certificate of Deposit

CP – Commercial Paper

CRSP – Center for Research and Security Prices

EH – Expectations Hypothesis

E/P – Earnings-to-Price

GMM – Generalized Method of Moments

LIBOR – London Interbank Offer Rate

LIFFE – London International Financial Futures and Options Exchange

LM – Lagrange Multiplier

NASDAQ – National Association of Security Dealers Automated Quotation

OLS – Ordinary Least Squares

OTC – Over-the-Counter

PH – Preferred Habitat

SIMEX – Singapore International Monetary Exchange

TIFFE – Tokyo International Financial Futures Exchange

VAR – Vector Autoregression

PART 1. PREFERRED HABITAT FOR LIQUIDITY IN INTERNATIONAL SHORT-TERM INTEREST RATES

1.1 Introduction and Literature Review

Pricing or valuing various securities is the cornerstone of the discipline of finance. Naturally, the concept of efficient markets is central to finance theory since market efficiency implies an environment in which assets may be priced correctly, according to their fundamental values. Any unusual patterns in security prices and/or volatilities will sooner or later be identified and draw attention of the academic world. These patterns may be associated with various characteristics of securities (e.g., size), markets (e.g., trading rules), general economic conditions, etc. Regularities that have not been rationally explained are labeled anomalies. The following literature review concentrates on the patterns associated with calendar times, particularly on those related to month-, quarter-, and year-ends. It does not extend to cover day-of-the-week and intraday regularities in security returns in order to maintain the desired focus of the study.

1.1.2 Empirical Studies

Time-based anomalies have been detected in different securities around the turn of a calendar period such as a year, quarter, month, week, and even a day (the extensive market microstructure literature deals with intraday data). This study focuses on the behavior of money market instruments at the turn of the quarter and year. Behavior of various securities around month-ends, quarter-ends, and year-ends has attracted a lot of academic attention in the past two and a half decades. For example, Reinganum (1981) documents that small firms outperform large firms on a risk-adjusted basis at the turn of the year; Jordan and Jordan

(1991) find a turn of the year effect in bonds; and Park and Reinganum (1986) report that T-bills maturing just before the end of the month, especially just before the end of the year, have lower yields than adjacent maturity T-bills.

Banz (1981) studies the relationship between market capitalization and common stock returns. He finds that, for the period 1936 through 1975, common stocks of small firms had, on average, higher risk-adjusted returns than common stocks of large firms. The identified size effect is not linear in market value: the difference is most striking between returns of very small firms and large firms, while it is much less significant between average sized and large firms. Risk has been controlled for by grouping securities into five beta quintiles and analyzing the size effects within these portfolios. The results do not reveal whether size itself drives excess returns or size is just correlated with some other factors that could explain the phenomenon. One conjecture expressed by Banz is that lack of information about small firms leads to limited diversification and therefore to higher returns on less desirable small stocks. He also notes that the size effect implies a misspecification of the traditional CAPM, since it improves the explanatory power of the model.

Reinganum (1981) constructs portfolios of common stocks based on E/P ratios and market values. He finds that low E/P portfolios and low market value portfolios earn abnormally high returns on a persistent basis. When studying the two factors together, Reinganum finds that the abnormal returns of high E/P portfolios become insignificant after controlling for size, while the abnormal returns of low market value portfolios remain significant after controlling for the E/P ratio. Thus the identified phenomenon in the behavior of returns is likely to be more closely associated with (or driven by) market values than E/P ratios. Reinganum, like Banz (1981), notes that his findings can be interpreted either as a sign

of a misspecification of the CAPM or evidence of market inefficiency, and tends to favor the former interpretation because persistency of the results over time is unlikely to be a result of market inefficiency.

Keim (1983) focuses on empirical month-by-month relations between market value and abnormal returns. He shows that, for the period 1963 through 1979, the major portion of size-related abnormal returns is driven by January abnormal returns. While January abnormal returns are responsible for about fifty percent of the “size effect”, more than a half of the January premium is attributable to the first five trading days in a year. The relationship between market value and returns is negative and more pronounced in January than in any other month, even in years when large firms outperform small firms on a risk-adjusted basis.

Taking into account the critique of Roll (1981), Scholes and Williams (1977) and Dimson (1979), all of whom argue that small stocks’ betas may be underestimated due to infrequent trading, Keim uses three different beta estimates (OLS beta, Scholes-Williams beta, and Dimson beta) but finds that the size effect is not sensitive to different estimators of beta. This is consistent with Reinganum (1982), who reports that, although small stocks’ OLS betas tend to be underestimated, confirming Roll’s (1981) conjecture, this does not eliminate a pronounced negative relation between abnormal returns and firm size.

Keim offers (but does not test) two possible explanation of the January effect: the tax-loss selling and the information hypothesis. Tax-loss selling refers to investors selling loser stocks in the end of the year in order to reduce capital gains tax; the information hypothesis states that the time after the fiscal year-end (December 31 for most firms) is a period of increased uncertainty due to impending release of important information.

Chan and Chen (1988), however, show that the size effect is insignificant when betas are estimated over the long run. The "size effect" is observed if only five years of data are used to estimate betas; it disappears when data for a long period of time are used to estimate betas. This finding implies that the January effect is a return seasonality rather than a result of a misspecification of the CAPM.

Lamoureux and Sanger (1989) analyze returns on NASDAQ stocks over the period 1973-1985. OTC-traded firms are on average much smaller than exchange-listed companies; also, a portfolio of smallest NASDAQ firms will not predominantly consist of stocks whose value has declined recently, contrary to the portfolio of exchange-listed smallest firms. A monotonic inverse relation between firm size and January excess returns is found, which implies that the January effect in stocks cannot be explained by institutional features of exchanges. Transactions costs are large enough to preclude investors from profitably exploiting the turn-of-the-year seasonality in small OTC-traded stocks.

Jordan and Jordan (1991) test for seasonal patterns in corporate bond returns and compare them to the seasonal patterns previously identified in equities. They identify several seasonalities in the bond markets using the Dow Jones Composite Bond Average. Corporate bonds exhibit significant turn-of-the-year and week-of-the-month effects (the highest mean return occurs in the second week of the month; week 4 has the lowest return). A January effect also exists but it is less strong than in equities. The turn-of-the-year effect being significant in both stocks and bonds lends credence to an explanation of the phenomenon that is general to fixed-income and equity markets.

1.1.3 Explanations of Month-, Quarter-, and Year-end Regularities

Three different, but not mutually exclusive hypotheses have been dominant in the literature attempting to explain these phenomena: the tax-loss selling (e.g., Roll (1983)), the window dressing (Haugen and Lakonishok (1988)), and the preferred habitat (Ogden (1987)). Empirical studies provide some support for all three of them: tax-loss selling (Jones et al. (1991), Griffiths and White (1993)), window dressing (Lakonishok et al. (1991), Allen and Saunders (1992)), and preferred habitat (Ogden (1987, 1990), Griffiths and Winters (1997, 2004)) in various markets.

1.1.3.1 Tax-loss Selling

The tax-loss selling hypothesis states that investors who want to realize capital losses in the current tax year dispose of stocks that have recently declined in value, thus putting downward pressure on prices.

Roll (1983) shows that over the period 1963 through 1981, the five largest daily mean return differences between an equally weighted and a value-weighted CRSP indexes occurred on five consecutive days: the last trading day of December and the first four trading days of January. About 37% of the return differential is attributable to these five days, and 67% of it occurs on the first 20 trading days of January, plus the last trading day in December. This implies that small firms outperform large firms over the course of the year, and the bulk of the “size effect” occurs on the five consecutive days commencing with the last trading day of December. Roll goes on to show that there is a significant negative relation between the turn-of-the-year return and the preceding year return on the stock. This is consistent with tax-loss selling at the end of the calendar year followed by a rebound in

prices when the selling pressure subsides after the turn of the year. Small stocks are more likely to experience substantial negative returns over some period of time due to the mere fact that they are much more volatile than stocks of large firms. A closer look at transaction costs and liquidity of such stocks may explain why the turn-of-the-year effect is not arbitraged away.

Ritter (1988) notes that, although the tax-loss selling accounts for a large part of abnormal January returns of small stocks, it does not explain the high January returns for small stocks that have not declined in value thus presenting investors with an opportunity to realize capital losses. He proposes the parking-the-proceeds hypothesis that focuses on buying and selling behavior of individuals around the turn of the year: some of the proceeds from December sales may not be reinvested immediately, but “parked” until January. The hypothesis requires that individual investors hold more of low-priced, low-capitalization stocks (Blume and Friend (1983)) whose prices are more dependent on buying and selling pressure (Lakonishok and Smidt (1984)), and do not reinvest the proceeds from tax-loss sales immediately in the same or different stocks. Using the buy/sell ratio for individual investors at Merrill Lynch, Ritter demonstrates that December’s net selling transforms into net buying at the turn of the year, which is consistent with the proposed parking-the-proceeds hypothesis. This framework is capable of explaining why small stocks do well at the turn of the year; however, it cannot explain why small stocks outperform large stocks over the course of the year.

Jones, Lee, and Apenbrink (1991) test whether the turn-of-the-year effect was present in common equities before the introduction of the U.S. personal income tax in 1917 under the War Revenue Act. Using an equally weighted Cowles Industrial Index, they find that the

turn-of-the year returns on this index are not significantly different from the returns for the Dow Jones Industrial Average in the pretax period; however, the difference becomes significant after the introduction of personal income taxes. Also, stocks with relatively high tax-loss potential earned a significantly higher turn-of-the-year returns in the post-tax period. Similar results hold for low-price stocks versus high-price stocks. The findings relate the January effect to the existence of income taxes; however, they do not provide a complete solution for the size effect.

Griffiths and White (1993) compare the turn-of-the-year effects in the U.S. and Canadian markets using intraday data. The tax year in Canada effectively ends five days prior to the calendar year-end, inducing investors who want to realize tax losses to trade at least five days before the calendar year-end. By analyzing whether transactions occur at bid or ask prices, Griffiths and White conclude that there is significant selling pressure before the tax year-end (most transactions occur at the bid price), with a shift to the buying pressure after the tax year-end when most transactions occur at the ask price in both countries. The findings provide further support for the tax-loss selling hypothesis and are consistent with the parking-the-proceeds hypothesis of Ritter (1988).

1.1.3.2 Window Dressing

Window dressing is essentially an attempt by disclosing entities to disguise the composition of their holdings by investing (disinvesting) disproportionately in some assets on or right before disclosure dates. This activity is temporary in nature, as window dressers are expected to reverse their positions after a disclosure date. Lakonishok et al. (1991) investigate whether pension fund managers are involved in window dressing around quarter-

ends and year-ends. Selling losers or buying winners before disclosure dates is a sign of window dressing. The results suggest that funds predominantly use the contrarian strategy (buy losers and sell winners) except they get rid of mistakes. Dumping loser stocks is more typical of smaller funds, probably because it is too hard to fool sophisticated investors of larger funds by window dressing. Overall, the findings are indicative of modest window dressing by some funds.

Allen and Saunders (1992) analyze behavior of bank assets around (quarterly) disclosure dates, and interpret increases in the level of some asset categories (e.g., Fed funds) as a clear sign of window dressing by banks. Turn-of-the-quarter increases in some liabilities (e.g., transaction deposits) are viewed by the authors as possible evidence of window dressing by bank customers.

Musto (1997) finds that commercial paper maturing in the next calendar year sells at an extra discount while T-bills do not. For example, in the data used, the 30-day commercial paper rate is 47 basis points higher just before the year-end than just after. The discount is even bigger for paper with higher default risk. He also finds quarter-end discounts, but no month-end discounts. The patterns for quarter-ends are similar to the year-end, although on a smaller scale. These results are interpreted by Musto as a sign of window dressing by intermediaries who do not want to have commercial paper on books on the year-end disclosure date, December 31. The argument is that intermediaries switch to riskless securities in an attempt of “flight-from-risk” across the year-end. By doing so, the intermediaries underrepresent the riskiness of their portfolios to investors on disclosure dates.

Musto (1999) studies weekly allocations of money market funds among various categories of securities. He notes that retail funds tend to hold larger proportions of assets in

riskless government securities around disclosure dates (every six months). On average, retail funds hold 0.3% of portfolio value more in government securities during disclosures than on other dates, with the difference being statistically significant. These reallocations are interpreted as a sign of window dressing. Because 0.3% is a small change to materially affect investors' decision, it is likely that the results are driven by few funds tilting their portfolios more toward riskless securities around disclosure dates. Partitioning the sample of retail funds into size categories does not yield any significant results, while categorizing funds with respect to their recent performance reveals that relatively poor performers tend to window dress more. The reason for such funds to window dress is perhaps that investors do not know whether low money-fund returns are due to poor performance or low risk. Still, given the small magnitude of the results, it is likely that only very few funds engage in window dressing around disclosures. Institutional funds (those with large minimum investments) do not tilt their portfolios toward government securities on disclosure dates. The offered explanation is that their clients can afford to buy weekly money fund reports, which is uneconomical for individual investors.

1.1.3.3 Preferred Habitat

Modigliani and Sutch (1966) developed the preferred habitat hypothesis to add to the explanation of the term structure of interest rates. The preferred habitat can be viewed as a variant of the market segmentation hypothesis of the term structure. The proponents of markets segmentation argue that investors have clearly defined preferences for instruments of certain maturities and stick to those maturities due to risk aversion. Institutional and regulatory constraints may determine such preferences. The preferred habitat hypothesis is

more flexible in that it states that investors can be induced into different maturities given large enough differences in yields.

Park and Reinganum (1986) document that T-bills maturing at the ends of calendar months have significantly lower yields than T-bills maturing at the beginning of calendar months. This difference is especially large at the turn of the year. Although the authors suggest some potential explanations, they declare such behavior of T-bill yields puzzling.

Ogden (1987) relates the phenomenon identified by Park and Reinganum to the aggregate preferred habitat for lenders at the ends of calendar months. He points out that about 70 percent of the interest and principal payments on corporate debt and 38 percent of dividend payments on common stocks are made either on the last or on the first business day of the month. Other contractually scheduled payments such as wages and salaries occur on these dates thus suggesting that an aggregate preferred habitat may influence yields of money market securities maturing at the turn-of-the-month. The analysis of T-bill yields conducted by Ogden supports the preferred habitat hypothesis: a preference by corporations and other institutions for the last bill maturing in a month results in lower yields for such bills relative to bills maturing a week later. Moreover, the interest rate differential between the first bill maturing in a month and the last bill maturing in the previous month widens as the latter matures closer to the end of the month.

Simon (1991) examines cash management T-bills which represent unexpected additional supplies of outstanding T-bills, and shows that the segmentation in the T-bill market is not limited to month-end preferred habitats.[1] He finds that such announcements

[1] In order to study only unexpected cash management issues, Simon excluded bills auctioned in place of regularly scheduled actions delayed due to debt ceiling constraints.

generally lead to an increase in cash management bill yields but do not influence adjacent maturity bill yields. During the sample period (1980 through 1988), the spread between cash management bill yields and the average of the two adjacent T-bill yields had increased on average by a statistically and economically significant 35 basis points from one day before to nine days after the announcement. The changes in the spreads are more pronounced for short maturity cash management bills, which implies that the segmentation is greater at the short end of the market. The unwillingness of market participants to reshuffle their portfolios to take advantage of arbitrage opportunities apparently reflects preferred habitats that extend beyond the end of the month.

Simon (1994) presents additional evidence on segmentation in the T-bill market by examining the yield differences between 13-week and 12-week T-bills. Some 12-week and 13-week issues represent additional supplies of previously auctioned 26- or 52-week bills. Controlling for factors that were previously identified to have influence on T-bill yield differentials (the slope of the yield curve, month-end preferred habitats, and cash management issues), Simon finds that the yield spreads between 13-week and 12-week bills are affected by differences in supply of these bills. In particular, yields on bills that originate as 52-week bills are about 4 basis points higher than they otherwise would be. Investors placing new funds in such issues would have an extra return of $100 on each $1 million. The fact that these yield differences are not arbitraged by market participants is interpreted as evidence of further segmentation of the T-bill market.

Griffiths and Winters (1997) use daily data on term repo rates to analyze the year-end regularity in the money markets. They start by reporting that one-, two-, and three-week repo spreads decline significantly in the last two trading days of the year, and the declines

continue through the first four trading days of the new year. This result rejects window dressing at the turn of the year. The direction of the spread changes contradicts idle cash window dressing (that is, an attempt to show little cash on books on disclosure dates to appear more efficient), while the timing of these changes is inconsistent with either the idle cash or the flight from risk variant of window dressing. Griffiths and Winters proceed to show that these turn-of-the-year declines in spreads are actually preceded by significant increases. The term repo spreads over the 91-day T-bill yield jump when the maturity of the instrument spans the end of the year. [2] For example, the two-week repo spreads over the T-bills are significantly higher on days -10 through -1 relative to the year-end than on other trading days. Trading day -10 represents the first day when maturity of the two-week repo spans the year-end. Similar patterns are observed for one-week, two-week, and one-month repos, but not in overnight or longer term repos. The findings cannot be consistent with year-end window dressing by institutions because of the timing and direction of spread changes, but are in line with the year-end preferred habitat for liquidity (Ogden (1987)). That is, investors prefer money-market securities that mature prior to the calendar year-end, so that they can fulfill their year-end obligations. Securities that mature across the end of the year will sell at a deeper discount due to lower investor demand for these maturities. Griffiths and Winters believe that this turn-of-the-year liquidity premium reflects preference for cash at the year-end. The demand for the securities should rebound after year-end obligations are paid and some cash is left over. This does not have to occur after the calendar year-end but can actually start one or several business days prior to that. This is exactly the pattern identified

[2] The 91-day T-bill was chosen because neither Musto (1997) nor Griffiths and Winters (1997) found any year-end regularity in this data series.

by Griffiths and Winters: the yields start falling back to normal levels before the calendar year-end, on trading day -1; the decline continues in the first few trading days of the new year.

Griffiths and Winters (2004) revisit the commercial paper market and find evidence that enables them to reject turn-of-the-year risk-shifting window dressing proposed by Musto (1997). Their results are consistent with year-end preferred habitat for liquidity due to year-end cash-flow obligations. This is similar to the findings of Griffiths and Winters (1997) for term repos. Griffiths and Winters (2004) study yield changes at the turn of the year more thoroughly and find that the decline in yields occurs prior to the calendar year-end. In particular, they find a significant rate decrease in a one-month CP over the last two trading days in December. Under the risk-shifting window dressing hypothesis, however, the yield increase must be sustained throughout the last trading day of the year because the hypothesis is that window dressers do not want to disclose risky positions in their year-end reports and thus must remain out of the risky security until after the year-end reporting date (12/31). Alternatively, under the preferred habitat hypothesis, the yield change patterns do not have to be perfectly aligned with the calendar year-end. Extending the analysis of rate changes to private-issue money market instruments with the original maturity of one month (banker's acceptances, negotiable CDs, Eurodollar deposits, and LIBOR), Griffiths and Winters find similar yield change patterns: the rates increase significantly in the last two trading days of November (this is just before a one-month instrument begins to mature across the year-end), followed by a return to "normal" levels in the last two trading days of December.

Griffiths and Winters (2003) extend the analysis of Griffiths and Winters (2004) by using spreads between different risk classes to determine if the price of risk increases at the end of

the year. They collect daily rates on 7-day, 15-day, and 30-day non-financial commercial paper for the two risk categories, AA and A2/P2, for the period 7/1/97 through 6/30/02. The year-end patterns for each maturity are consistent with those identified by Griffiths and Winters (2004). In addition, the spread between AA and A2/P2 commercial paper yields behaves similarly to the rates themselves. That is, the spread widens when maturities of the instruments start to span the turn of the year, and it returns to normal levels prior to the calendar year-end. For example, the 15-day average risk spread increases from 29 basis points on trading day -11 to 78 basis points on trading day -9 relative to the year-end. It returns to 27 basis points on trading day -1. These results suggest that there is an increased price for risk at the turn of the year, with the timing of the spread changes being consistent with preferred habitat for liquidity but not with risk-shifting window dressing.

Ogden (1990) notes that a standardization of the payments system in the U.S. may help explain the monthly effect identified by Ariel (1987). Ariel documents that virtually all of the cumulative returns on value weighted and equally weighted stock indexes, over the period from 1963 through 1981, are realized on the ten trading days starting with the last trading day of the month and ending with the ninth trading day of the following month. Lakonishok and Smidt (1988) find that for the Dow Jones Industrial Average, significant daily returns are realized only on four consecutive trading days starting with the last trading day of the month. Ogden's proposed explanation is that investors, many of whom have substantial cash receipts at the turn of the month, will increase their demand for stocks at the time. He also hypothesizes that expected liquid profits at the turn of the month are inversely related with the stringency of monetary policy (measured by the spread of the Fed funds rate over the one-month T-bill rate). This hypothesis is supported by Ogden's analysis of the CRSP stock

index returns (both value-weighted and equally weighted) for the period from 1969 through 1986.

1.1.4 Conclusion

The purpose of this study is to further develop the investigation of time-based anomalies in short-term debt markets. In particular, the main goal is to determine whether year-end and quarter-end effects in short-term interest rates exist for debt instruments denominated in currencies other than the U.S. dollar. Since short-term rates reflect yields of money market instruments which have low default and interest rate risk, tax-loss selling is not likely to be a plausible explanation for the turn-of-the-year effect in these markets. Therefore we are left with two major hypotheses: window dressing by market participants and quarter-end/year-end preferred habitat for liquidity. Although these hypotheses are not mutually exclusive, one of the explanations may dominate in any particular market. If the flight-from-risk variant of window dressing by intermediaries and other money market participants dominates around quarter- and year-ends, then short-term interest rates are expected to increase at these times and to return to normal levels after the end of the calendar quarter/year. If idle-cash window dressing is the driving force of the turn-of-the-quarter effect, short-term interest rates should decrease before and return to normal levels after the end of the calendar period. Under quarter-end/year-end preferred habitat for liquidity, interest rates for short-term securities would increase at the time when maturity of the instrument starts to span the end of the quarter or year, and return to normal levels after the quarter- or year-end cash obligations have been fulfilled. If many market participants have financial commitments on or immediately before the end of the quarter, they would prefer to invest in money-market

securities that mature just prior to the quarter-end. To induce these investors to purchase securities maturing in the next quarter, a premium would have to be offered. By the analogy with Griffiths and Winters (1997), such a premium will be labeled the quarter-end liquidity premium, and the phenomenon will be referred to as the quarter-end preferred habitat for liquidity. The premium may disappear before the end of the quarter/year, as long as cash obligations are fulfilled and free funds are available to be reinvested. In fact, evidence presented in Griffiths and Winters (1997, 2003, 2004) demonstrates that this is the pattern exhibited by the yields of the U.S. money market instruments.

1.2 Data and Methods

1.2.1 Data

The data are daily short-term London Interbank Offer Rates (LIBOR) for different currencies and maturities; they were obtained from the British Bankers' Association (BBA) website. The time periods vary for different currencies and maturities.

BBA LIBOR is the primary benchmark for short term interest rates globally. It is used as the basis for settlement of interest rate contracts on many of the world's major futures and options exchanges including London International Financial Futures and Options Exchange (LIFFE), Deutsche Term Burse, Chicago Mercantile Exchange, Chicago Board of Trade, Singapore International Monetary Exchange (SIMEX), and Tokyo International Financial Futures Exchange (TIFFE), as well as most over the counter (OTC) and lending transactions. It is compiled by the BBA in conjunction with Moneyline Telerate and released to the market shortly after 11:00AM London time each trading day.

The following is an excerpt from the definition of the BBA LIBOR fixing:

BBA LIBOR is the BBA fixing of the London Inter-Bank Offered Rate. It is based on offered inter-bank deposit rates contributed in accordance with the Instructions to BBA LIBOR Contributor Banks. The BBA consults on the BBA LIBOR rate fixing process with the BBA LIBOR Steering Group. The BBA LIBOR Steering Group comprises leading market practitioners active in the inter-bank money markets in London.

Contributor Panels shall comprise at least eight Contributor Banks.[3] *Contributor Panels will broadly reflect the balance of activity in the inter-bank deposit market. Individual Contributor Banks are selected by the BBA's FX & Money Markets Advisory Panel after private nomination and discussions with the Steering Group, on the basis of reputation, scale of activity in the London market and perceived expertise in the currency concerned, and*

[3] Of course, contributor panels are different for different currencies. As of July 23, 2002, for example, the contributor panels for Australian Dollar, New Zealand Dollar, and Danish Krone were comprised of eight banks, for Canadian Dollar and Swiss Frank – of 12 banks, and for Euro, Pound Sterling, Japanese Yen, and U.S. dollar – of 16 banks.

giving due consideration to credit standing. The BBA, in consultation with the BBA LIBOR Steering Group, will review the composition of the Contributor Panels at least annually.

Contributed rates will be ranked in order and only the middle two quartiles averaged arithmetically.[4] *Such average rate will be the BBA LIBOR Fixing for that particular currency, maturity and fixing date. Individual Contributor Panel Bank rates will be released shortly after publication of the average rate. In the event that it is not possible to conduct the BBA LIBOR Fixing in the usual way, the BBA, in consultation with Contributor Banks, the BBA LIBOR Steering Group and other market practitioners, will use its best efforts to set a substitute rate. This will be the BBA LIBOR Fixing for the currency, maturity and fixing date in question. Such substitute fixing will be communicated to the market in a timely fashion. An individual BBA LIBOR Contributor Panel Bank has to contribute the rate at which it could borrow funds, were it to do so by asking for and then accepting inter-bank offers in reasonable market size just prior to 11a.m.*

BBA LIBOR's London base is very significant: well over 20% of all international bank lending and more than 30% of all foreign exchange transactions take place through the offices of banks in London; it represents a unique snapshot of competitive funding costs. Close to 500 banks are represented in London, along with many other major financial institutions actively trading in the Euromarkets, which are based primarily in London. In addition, no reserve requirements are applied in London. Currently, BBA LIBOR fixings are provided in nine international currencies: Pound Sterling (GBP), U.S. Dollar (USD), Japanese Yen (JPY), Swiss Franc (SF), Canadian Dollar (CAD), Australian Dollar (AUD), Euro, Danish Krone (DK), and New Zealand Dollar (NZD).[5] LIBOR rates are fixed for each of these currencies in 15 maturities: overnight, one week, two weeks, and one month through twelve months. The BBA commenced fixing overnight rates (for GBP, EUR, CAD and USD) and spot/next rates (for AUD, SF and JPY) on January 2, 2001. The two-week LIBOR has also been fixed since the beginning of 2001, while the one-week LIBOR has been reported since the beginning of 1998. The euro-in zone currency LIBOR (those eventually replaced by

[4] This implies that daily LIBOR are not influenced by large liquidity squeezes/excesses in one bank.
[5] LIBOR fixings for Danish Krone and New Zealand Dollar commenced in July 2003. These currencies are not included in the analysis due to the short data period.

the Euro – German Mark, French Franc, Spanish Peseta, Italian Lira, Dutch Guilder, and Irish Punt) ceased to be fixed at the beginning of 1999. The first four out of these six currencies have been included in the analysis in this study. The primary focus of this work is on the shorter maturities, especially one week and one month, as it has been found in the past that these maturities exhibit the most noticeable year-end and quarter-end effects.

Table 1. Data periods for LIBOR currencies and maturities

		Maturities			
Currencies	**Abbrevia-tions**	**One day**	**One week**	**Two weeks**	**One-twelve months**
U.S. Dollar	USD	1/01-3/04	1/98-3/04	1/01-3/04	1/87-3/04
British Pound	GBP	1/01-3/04	1/98-3/04	1/01-3/04	1/87-3/04
Australian Dollar	AUD	1/01-3/04	1/98-3/04	1/01-3/04	1/89-3/04
Swiss Franc	SF	1/01-3/04	1/98-3/04	1/01-3/04	1/89-3/04
Euro	-	1/01-3/04	1/98-3/04	1/01-3/04	1/89-3/04
Japanese Yen	JPY	1/01-3/04	1/98-3/04	1/01-3/04	1/89-3/04
Canadian Dollar	CAD	1/01-3/04	1/98-3/04	1/01-3/04	7/90-3/04
German Mark	DM	N/A	1/98-12/98	N/A	1/87-12/98
French Franc	FF	N/A	1/98-12/98	N/A	1/89-12/98
Spanish Peseta	SP	N/A	1/98-12/98	N/A	7/90-12/98
Italian Lira	ITL	N/A	1/98-12/98	N/A	7/90-12/98

Table 1 summarizes the data periods for different LIBOR currencies and maturities. It also contains currency abbreviations used throughout the paper.[6]

Overall, BBA LIBOR probably represents the widest possible range of maturities for money-market securities. Table 2 presents descriptive statistics for LIBOR for monthly maturities of the 11 currencies included in the analysis.

Interest rates in all of the countries included in the analysis were much lower in the second half of the data periods than in the first several years. A glance at Table 2 confirms this statement: the difference between the average one-month LIBOR (available since

[6] Although the Euro came into existence in the beginning of 1998 for interbank payments and in the beginning of 2002 for cash payments, BBA LIBOR fixings have commenced for its predecessor, European Currency Unit (ECU), in January 1989. Thus, when we refer to the Euro, we also refer to its predecessor.

January 1987, January 1989, or July 1990, depending on the currency) and the average one-day LIBOR (available since 2001) is very dramatic varying from the low of 89 basis points for the Euro to the high of 351 basis points for the British Pound.

The term structures for the maturities of one through twelve months have been upward sloping for the first eight currencies in Tables 1 and 2. The yield curves increase monotonically for all of these currencies except the Euro which has a flat interval for maturities between three through nine months. The yield curve for the USD has the steepest slope, while that for Euro is closest to flat among the first eight currencies. In the case of the French Franc, the term structure becomes inverted after maturity reaches three months, although it is closest to flat among all currencies. For two other Euro-in zone currencies, Spanish Peseta and Italian Lira, the yield curves are downward sloping on the short end, but they become virtually flat after maturity reaches six months.

Figures 1 through 3 plot the yield curves for maturities of one through twelve months for each of the currencies included in the analysis. Figure 1 depicts the yield curves for USD and GBP, Figure 2 contains the yield curves for AUD, SF, Euro, JPY, and CAD, and Figure 3 – for the currencies eventually replaced by the Euro (DM, FF, SP, and ITL). This division is not completely arbitrary: USD and GBP LIBOR are available for the period from 1/87 through 3/04, while the data period for the currencies in Figure 2 is 1/89 through 3/04 (with the exception of CAD for which LIBOR fixings commenced in 7/90). The third figure includes the currencies whose LIBOR data availability extends only through 12/98.

While yield curves in Figures 1 through 3 may have positive or negative slopes or be close to flat, they represent averages over the period of data availability. At any given point in time, yield curves may have looked quite differently from those in the figures.

Table 2. Descriptive statistics for different LIBOR currencies and maturities

	Maturities														
Currencies	**One day**	**One week**	**Two weeks**	**1 month**	**2 months**	**3 months**	**4 months**	**5 months**	**6 months**	**7 months**	**8 months**	**9 months**	**10 months**	**11 months**	**12 months**
U.S. Dollar															
-Observations	798	1579	822	4363	4363	4363	4363	4363	4363	4363	4363	4363	4363	4363	4363
-Mean	2.220	3.874	2.204	5.281	5.326	5.361	5.389	5.419	5.446	5.480	5.513	5.547	5.585	5.623	5.661
-St. deviation	1.431	2.049	1.378	2.204	2.215	2.224	2.226	2.230	2.232	2.234	2.234	2.236	2.233	2.230	2.228
British Pound															
- Observations	822	1579	822	4362	4362	4362	4362	4362	4362	4362	4362	4362	4362	4362	4362
-Mean	4.214	5.147	4.230	7.730	7.772	7.795	7.802	7.814	7.823	7.837	7.850	7.861	7.881	7.900	7.916
-St. deviation	0.858	1.279	0.691	3.305	3.298	3.284	3.266	3.245	3.226	3.205	3.184	3.162	3.143	3.124	3.106
Australian Dollar															
- Observations	821	1579	822	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855
-Mean	4.830	5.045	4.844	7.275	7.295	7.307	7.322	7.339	7.361	7.389	7.420	7.453	7.485	7.523	7.568
-St. deviation	0.401	0.514	0.399	3.722	3.730	3.718	3.702	3.688	3.681	3.671	3.663	3.654	3.644	3.638	3.680
Swiss Franc															
- Observations	822	1579	822	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855
-Mean	1.368	1.560	1.372	3.796	3.827	3.849	3.857	3.864	3.872	3.877	3.883	3.889	3.899	3.908	3.920
-St. deviation	1.231	1.085	1.216	2.847	2.839	2.826	2.805	2.787	2.774	2.753	2.729	2.707	2.692	2.671	2.656
Euro															
- Observations	828	1599	832	3875	3875	3875	3875	3875	3875	3875	3875	3875	3875	3875	3875
-Mean	3.246	3.467	3.239	5.958	5.972	5.984	5.984	5.988	5.990	5.988	5. 987	5.987	5.994	6.004	6.011
-St. deviation	0.958	0.881	0.922	2.862	2.865	2.870	2.863	2.862	2. 860	2.849	2.837	2.826	2.820	2.815	2.810
Japanese Yen															
- Observations	822	1579	822	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855	3855
-Mean	0.071	0.177	0.074	2.238	2.242	2.241	2.240	2.239	2.241	2.244	2.248	2.253	2.259	2.265	2.272
-St. deviation	0.111	0.283	0.097	2.668	2.666	2.658	2.645	2.636	2.630	2.621	2.612	2.604	2.601	2.595	2.594
Canadian Dollar															
- Observations	804	1579	822	3477	3477	3477	3477	3477	3477	3477	3477	3477	3477	3477	3477
-Mean	3.190	4.105	3.179	5.257	5.291	5.322	5.355	5.387	5.417	5.453	5.487	5.535	5.561	5.595	5.631
-St. deviation	0.973	1.218	0.941	2.304	2.284	2.264	2.243	2.228	2.215	2.202	2.193	2.178	2.178	2.165	2.157

	Maturities														
Currencies	**One day**	**One week**	**Two weeks**	**1 month**	**2 months**	**3 months**	**4 months**	**5 months**	**6 months**	**7 months**	**8 months**	**9 months**	**10 months**	**11 months**	**12 months**
German Mark															
- Observations		253		3037	3037	3037	3037	3037	3037	3037	3037	3037	3037	3037	3037
-Mean		3.478		5.789	5.817	5.843	5.858	5.872	5.882	5.884	5.886	5.892	5.902	5.914	5.925
-St. deviation		0.120		2.298	2.309	2.315	2.308	2.305	2.303	2.290	2.278	2.266	2.258	2.248	2.241
French Franc															
- Observations		253		2529	2529	2529	2529	2529	2529	2529	2529	2529	2529	2529	2529
-Mean		3.458		7.139	7.150	7.164	7.150	7.137	7.127	7.116	7.107	7.102	7.102	7.106	7.107
-St. deviation		0.080		2.849	2.824	2.809	2.776	2.748	2.727	2.707	2.689	2.676	2.666	2.656	2.648
Spanish Peseta															
- Observations		253		2151	2151	2151	2151	2151	2151	2151	2151	2151	2151	2151	2151
-Mean		4.407		9.570	9.512	9.467	9.432	9.406	9.380	9.374	9.373	9.369	9.373	9.380	9.386
-St. deviation		0.464		3.667	3.589	3.530	3.500	3.486	3.472	3.460	3.459	3.458	3.457	3.462	3.464
Italian Lira															
- Observations		252		2151	2151	2151	2151	2151	2151	2151	2151	2151	2151	2151	2151
-Mean		5.227		9.609	9.565	9.537	9.487	9.516	9.505	9.498	9.491	9.494	9.482	9.484	9.485
-St. deviation		0.821		2.919	2.814	2.774	2.871	2.805	2.822	2.841	2.861	2.871	2.895	2.910	2.924

Table 2 presents descriptive statistics for LIBOR of 11 currencies utilized in the analysis, which include the number of observations, means, and standard deviations for each of available maturities.

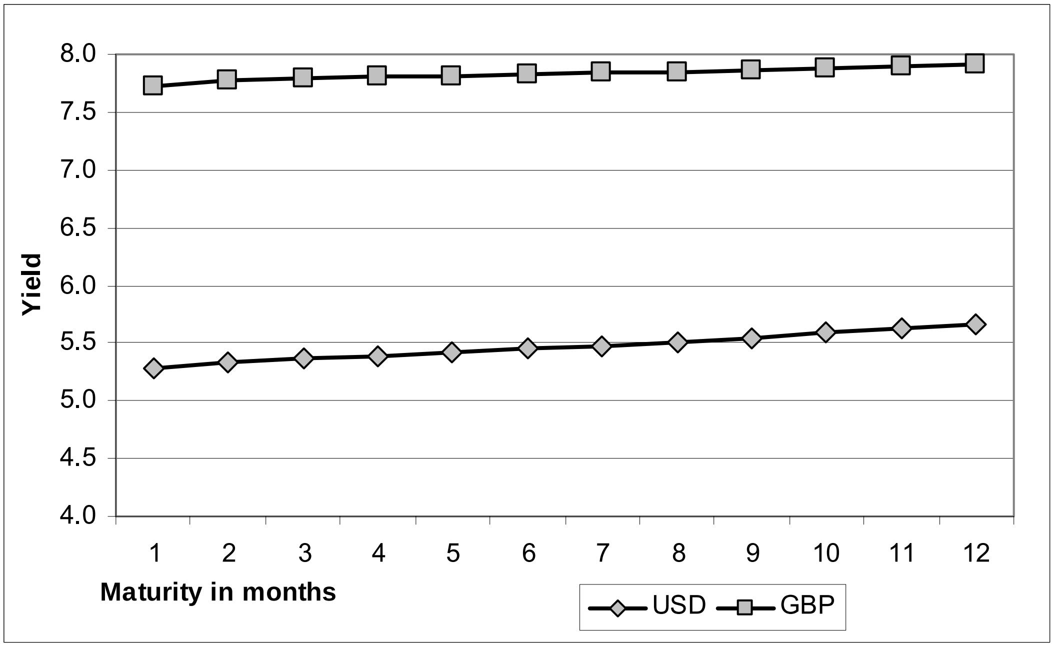

Figure 1. LIBOR yield curves for USD and GBP (January 1987 through March 2004) for maturities of one through twelve months.

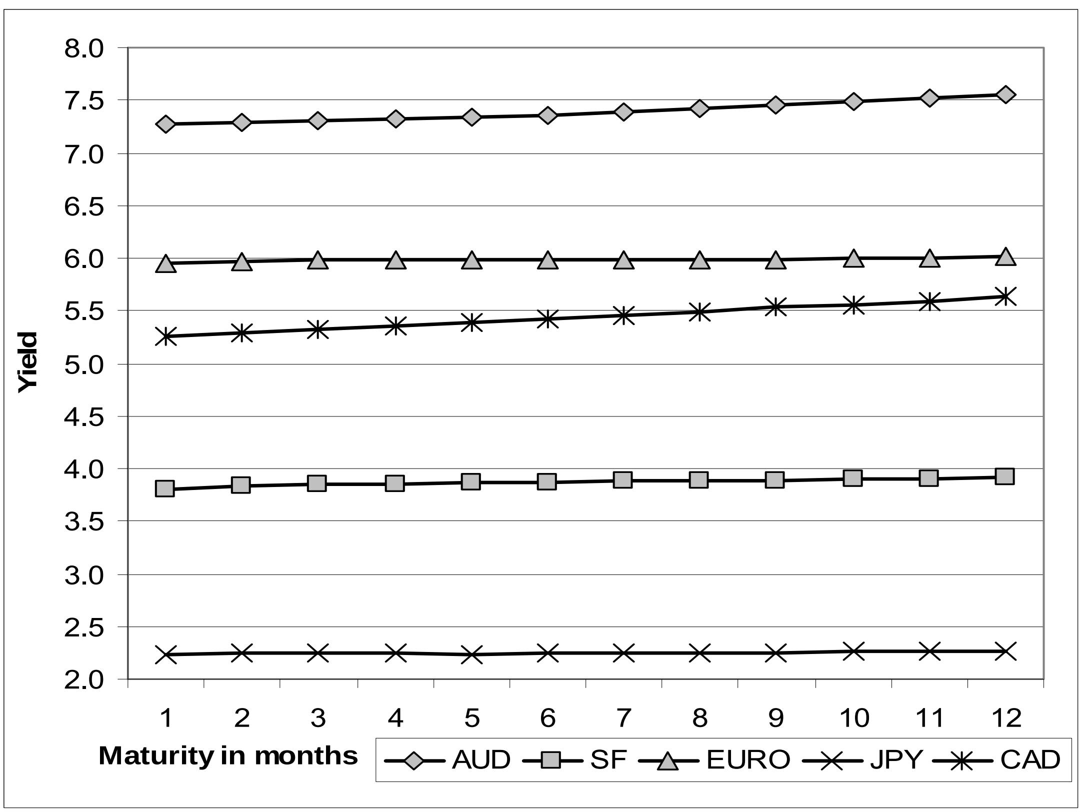

Figure 2. LIBOR yield curves for AUD, SF, Euro, JPY (January 1989 through March 2004), and CAD (July 1990 through March 2004) for maturities of one through twelve months.

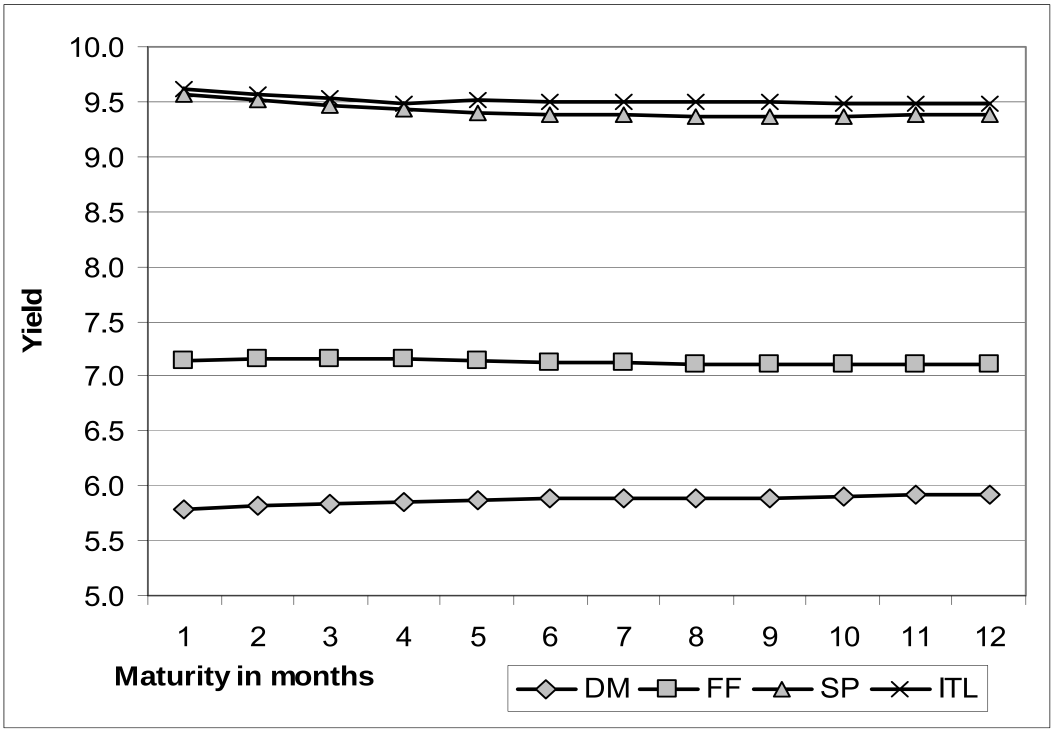

Figure 3. LIBOR yield curves for the Euro-in zone currencies: DM (January 1987 through December 1998), FF (January 1989 through December 1998), SP and ITL (July 1990 through December 1998) for maturities of one through twelve months.

Figures 4 through 6 plot volatility term structures for the same sets of currencies. Volatility curves for maturities from one through twelve months are downward sloping for GBP, SF, JPY, FF, CAD, and SP. For the first five of these currencies, standard deviations decrease monotonically as maturity increases. This is consistent with the observation of Fisher (1896) that shorter-term interest rates should fluctuate more than longer-term rates. The volatility curves for Euro and DM are upward sloping at the short end, but soon become downward sloping, while that for USD also has an upward slope at the short end but becomes virtually flat for longer maturities. In the case of ITL, the two highest standard deviations are at the opposite ends of the maturity specter: its volatility curve sags in the middle.

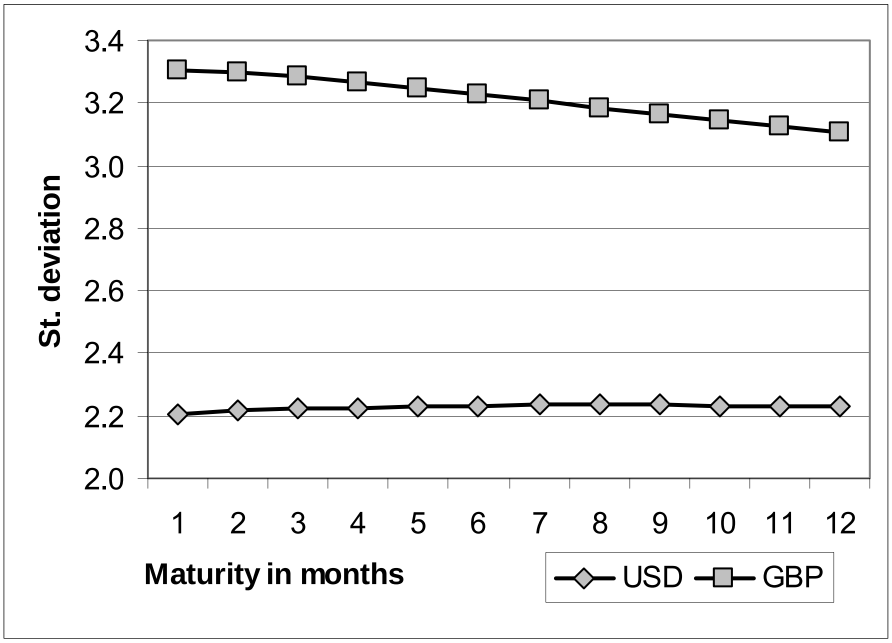

Figure 4. Volatility term structures for USD and GBP LIBOR maturities of one through twelve months (January 1987 through March 2004).

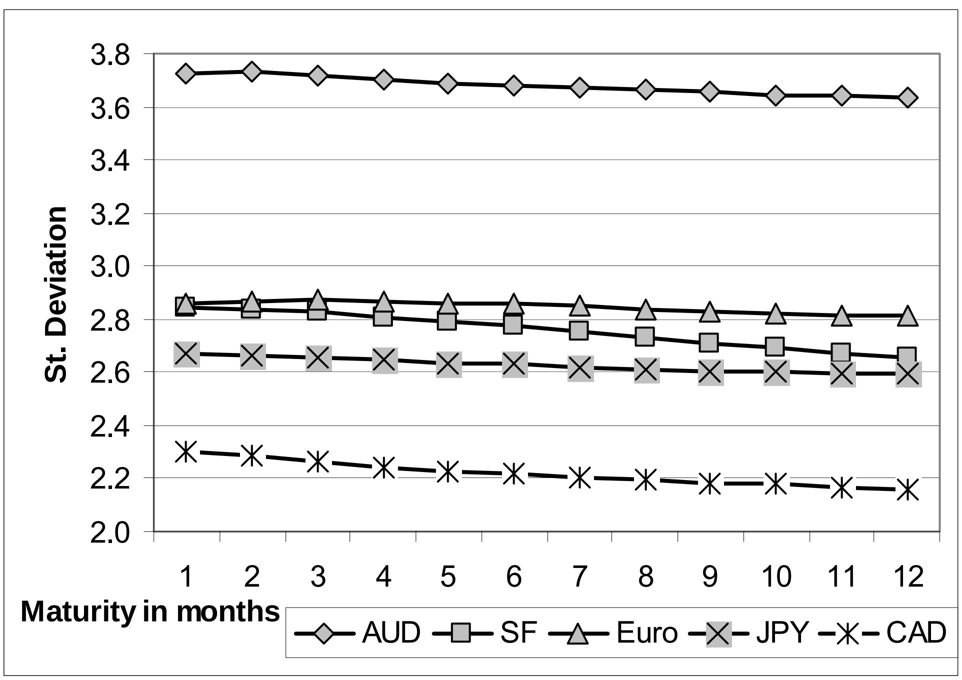

Figure 5. Volatility term structures for AUD, SF, Euro, JPY (January 1989 through March 2004), and CAD (July 1990 through March 2004) for maturities of one through twelve months.

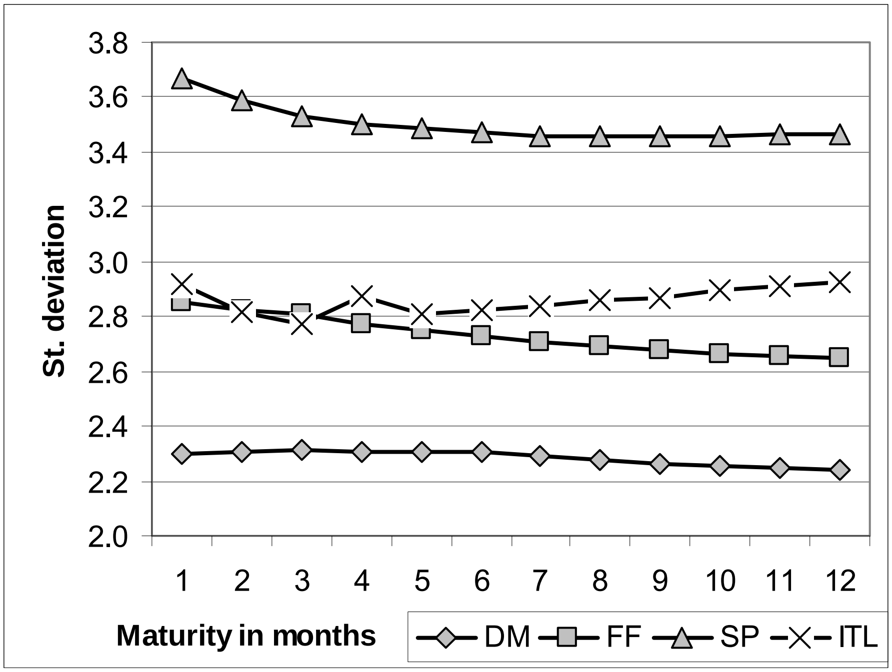

Figure 6. Volatility term structures for the Euro-in zone currencies:
DM (January 1987 through December 1998), FF (January 1989 through December 1998), SP and ITL (July 1990 through December 1998) for maturities of one through twelve months.

1.2.2 Methods

Jordan and Jordan (1991) compare equity and fixed-income securities seasonals using daily returns of bond and equity indexes. They start with the day-of-the week effect. They classify the first 10 trading days of each month as days 1, 2, 3, …, 10, and the last 10 trading days as -10, -9, …, -1. Each month thus has 20 trading days (days in excess of 20 were deleted, and in months with fewer than 20 trading days some days were counted twice). Trading days 1 through 5 constitute week 1, days 6 through 10 constitute week 2, days -10 to -6 are week 3, and days -5 to -1 are week 4. They test for the week-of-the-month effect in bonds and find some significant results. Finally, they test for the turn-of-the-month and turn-of-the-year. The turn of the month is defined as the first four trading days in each month plus the last trading day in the previous month. Jordan and Jordan use ANOVA to test whether returns are different across days, weeks, or months. They found that the January mean return on bonds is significantly higher than returns in other months. Half of the January mean return is attributable to non-turn-of-the-year days in January (they exclude turn-of-the-month days from the analysis to compare non-turn-of-the-month returns).

Musto (1997) compares the spreads between commercial paper of different maturities before and after one of them spans the end of the year. That way, he does not need a control for the general level of interest rates. He finds significant differences, but misses the exact timing of the year-end effect. He also conducts turn-of-the-quarter tests and finds some significance.

Griffiths and Winters (1997) focus on the year-end effect. They exclude all observations between January and November inclusive because they find no significant effects in those months, then run a regression with the dependent variable defined as a natural log of the

relative change in the spread between the repo rate and the three-month T-bill yield. The three-month T-bill yield was chosen because it exhibited no significant turn-of-the-year effect.

Griffiths and Winters (2004) create a turn-of-the-month dummy variable that covers trading days –2 through 4 in each month. For the turn of the year, this variable is split into two (the last two trading days of the year are covered by the dummy called YEND, the first four – by YBEG). Since they focus on a one-month commercial paper (CP), they also create two dummies for the turn of November (NOVEND – the last two days of November, and DECBEG – the first two days of December). They use the first difference of the CP rate as a DV and a change in the three-month T-bill rate as a control.

While Griffiths and Winters (2004) concentrate on the turn of the year, this study also focuses on "regular" quarter-ends. It is recognized that year-ends might be different from other quarter-ends. For example, Musto (1997) finds that the year-end effect in U.S. commercial paper is much larger in magnitude than yield changes related to other quarter-ends. Thus, two sets of dummy variables were created to make the distinction between year-ends and other quarter-ends. The following model is specified and estimated using OLS with White's (1980) adjustment for heteroskedasticity:

$$\Delta Rsp_t = \alpha_0 + \alpha_1(\Delta Rsp_{t-1}) + \alpha_2 BQCR + \alpha_3 AQCR + \alpha_4 BQEND + \alpha_5 AQEND + \alpha_6 BYCR + \alpha_7 AYCR + \alpha_8 BYEND + \alpha_9 AYEND + \varepsilon_t \quad (1)$$

where:

ΔRsp_t = the change in the relative spread between LIBOR of given currency and maturity and the three-month LIBOR for the same currency, $Rsp_t - Rsp_{t-1}$. The relative spread is defined as $R_t/3MR_t$, where R_t is the short-term LIBOR whose behavior is studied, and $3MR_t$ is the three-month LIBOR for the same currency, which serves as the base rate.

ΔRsp_{t-1} = the previous day change in the LIBOR of given currency and maturity, $Rsp_{t-1} - Rsp_{t-2}$

BQCR = a dummy variable that equals 1 on the two trading days before the maturity of the instrument starts to span the end of the quarter, *except* the fourth quarter, and 0 otherwise,

AQCR = a dummy variable that equals 1 on the two trading days after the maturity of the instrument starts to span the end of the quarter, excluding the fourth quarter, and 0 otherwise,

BQEND = a dummy variable that equals 1 on the last two trading days of the quarter, excluding the fourth quarter, and 0 otherwise,

AQEND = a dummy variable that equals 1 on the first two trading days of the quarter, excluding the fourth quarter.

BYCR, AYCR, BYEND, and AYEND are the dummy variables defined similarly to the previous four dummies; they are designed to isolate the turn of the fourth quarter. We will be referring to the ends of quarters one, two, and three as quarter-ends, and to the end of the forth quarter as the year-end hereafter.

For the sake of clarity, an example with actual dates would help visualize the timing of the turn of the quarter and the correspondence of particular days to the dummy variables. Table 3 below provides such an example under assumptions of the turn of the first quarter and LIBOR maturity of one week (seven calendar days). For ends of quarters one through three, there are always five trading days between the "issuance" of the one-week LIBOR and its maturity. At the turn of the year, however, only three trading days are spanned by the maturity of the instrument; this is because Christmas (December 25th) and Boxing Day (December 26th) are non-trading days in the U.K. It creates a situation where the two dummy variables, AQCR and BYEND, overlap in some years. This presents a methodological

problem and makes year-end results for the one-week LIBOR less reliable. This is not a problem for longer maturities, of course.

Table 3 also contains signs of the coefficients of the turn-of-the quarter dummies associated with each of the three competing hypotheses summarized in the literature review. This is done to help the reader visualize the dynamics of yield changes that would be consistent with each of the hypotheses.

Table 3. The turn-of-the-quarter dummy variables and their hypothesized signs

Variable/ Date	BQCR	AQCR	BQEND	AQEND	Hypothesized yield change signs		
					Flight from risk WD	Idle cash WD	Preferred habitat
3/23, Mon	1	0	0	0	?	?	+
3/24, Tue	1	0	0	0	?	?	+
3/25, Wed	0	1	0	0	+	–	?
3/26, Thu	0	1	0	0	+	–	?
3/27, Fri	0	0	0	0	?	?	?
3/28, Sat	0	0	0	0	?	?	?
3/29, Sun	0	0	0	0	?	?	?
3/30, Mon	0	0	1	0	?	?	–
3/31, Tue	0	0	1	0	?	?	–
4/1, Wed	0	0	0	1	–	+	?
4/2, Thu	0	0	0	1	–	+	?

Table 3 illustrates the timing of the dummy variables designed to test for the turn-of-the quarter effect and short-term yield changes consistent with each of the competing hypotheses. It has been developed for a hypothetical first quarter and one-week maturity.

The change in the relative spread was selected as a dependent variable. The measure of the relative spread was defined as $R_t/3MR_t$, where $3MR_t$ is the three-month LIBOR for a given currency.[7] The independent variables include the lagged change in the relative spread along with the turn-of-the-quarter and turn-of-the year dummies. The three-month LIBOR was chosen to be the base rate, or the control for the general level of interest rates for a given currency. Its choice was stipulated by the finding that the three-month LIBOR is the shortest

[7] Defining the relative spread as $(R_t - 3MR_t)/3MR_t$ does not make any difference because daily changes in these differently defined spreads are identical: $\Delta[(R_t - 3MR_t)/3MR_t] = \Delta[R_t/3MR_t]$

maturity that does not have quarter-end or year-end preferred habitats. It also reduces as much as possible the maturity mismatch between the rate being studied and the base rate.

Another possible choice for the dependent variable is a daily change in the interest rate, or simple first difference. This choice provides a very straightforward interpretation of the results: the coefficients of independent variables are basis points. However, one general concern that may be raised about this dependent variable is that it does not capture trends in rates or spreads. To address this concern, the change in the relative spread was selected as a dependent variable.[8]

One more remark is in order before proceeding to the results section. All studies cited above have dealt with private-issue U.S. securities such as commercial paper, repurchase agreements, bankers' acceptances, etc. The interbank (fed funds) market was avoided because it has well-documented regularities (see, for example, Cyree and Winters (2001)) that may make the analysis more difficult. While the data utilized in this study represent interbank loan rates, there is no reason to suspect that regularities similar to those in the U.S. federal funds market also exist in the London interbank market. Interbank rates may reflect the needs of bank customers, e.g., large fund withdrawals before a year-end may lead to higher interbank rates. Moreover, banks themselves may have quarter-end or year-end cash obligations such as interest payments on long-term debt and dividend payments to shareholders. If this is the case, we might see significant quarter-end and year-end effects consistent with the preferred habitat hypothesis.

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[8] Results are not qualitatively different if a simple first difference ($Rsp_t - Rsp_{t-1.}$), an absolute spread ($Rsp_t - 3MR_{t-1.}$), or a change in absolute spread is used as a dependent variable.

On the other hand, if banks are involved in idle-cash window dressing, the interbank rate is expected to decline before and rebound after disclosure dates as banks attempt to show less cash on books to appear more efficient. The flight-from-risk variant of window dressing would be feasible only if banks believe that interbank loans are perceived as risky by investors who read quarter-end and year-end disclosures. The use of LIBOR data provides a good opportunity for testing for quarter-end and year-end effects.

1.3 Empirical Results

1.3.1 One-week Results

Before proceeding to the discussion of regression output, a reader may find it useful to visualize the dynamics of quarter- and year-end behavior of short-term rates. These are graphically presented for one-week LIBOR of each of the seven currencies in Figures 7 through 20. The figures plot the average (absolute, not relative) spread for the period from January 1998 through March 2004 between the one-week and three-month LIBOR for each currency around quarter-ends. That is, the spread in basis points on trading day -10 relative to the end of the quarter has been averaged across all quarters for which both one-week and three-month rates are available. The same has been done to the spreads on trading days -9 through 10 relative to the quarter-end. Each quarter ends with day -1 and begins with day 1, that is, there is no day 0.

A pair of graphs describes turn-of-the-quarter and turn-of-the-year spread behavior for each currency starting with the U.S. Dollar. The first graph plots the following three series: 1) spreads around all quarter-ends and year-ends, 2) spreads around all quarter-ends and year-ends with the exception of Y2K, and 3) spreads around quarter-ends only. The second graph plots two average spread series that depict the spread behavior around year-ends only, one with and the other one without the observations around Y2K.

The patterns of spread changes presented in Figures 7 through 20 is consistent (in both timing and direction) with the quarter-end and year-end preferred habitat for liquidity in all seven currencies. These patterns are particularly pronounced in the U.S. Dollar, Swiss Franc, Euro, and Japanese Yen. For all seven currencies, the spread between one-week and three-

month LIBOR increases on days -7 through -5 relative to the quarter- or year-end, and returns to normal levels on the second to last trading day of the quarter or year. Regression analysis should provide a more definite conclusion regarding the statistical significance of these calendar effects.

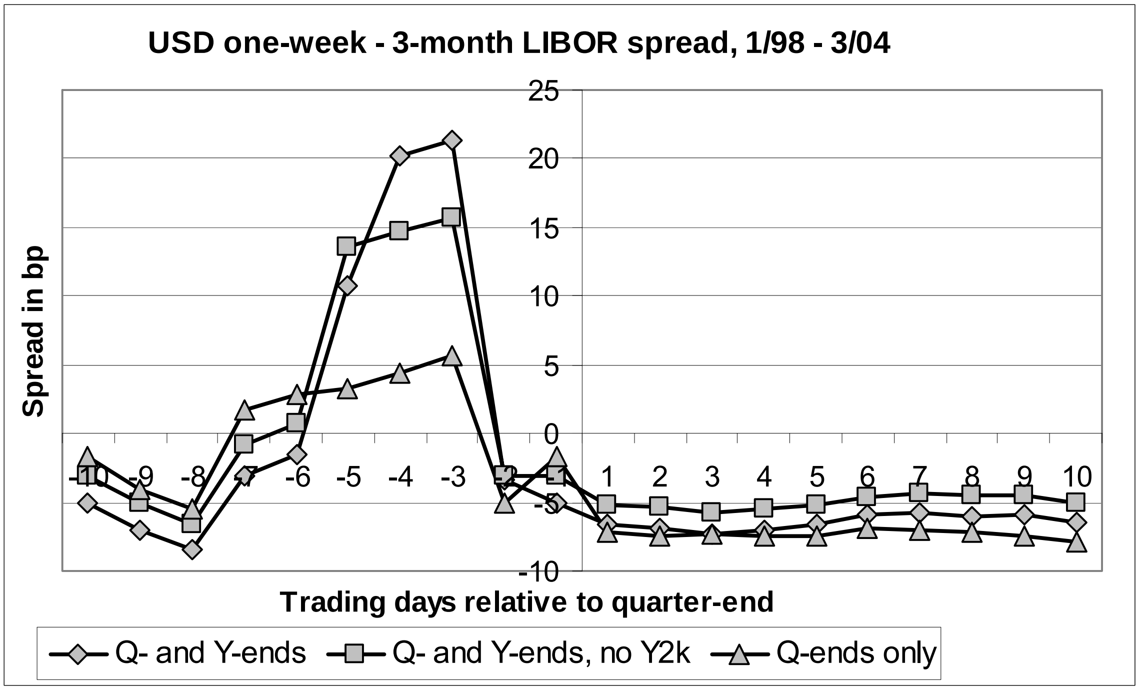

Figure 7. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for USD over the period January 1998 through March 2004.

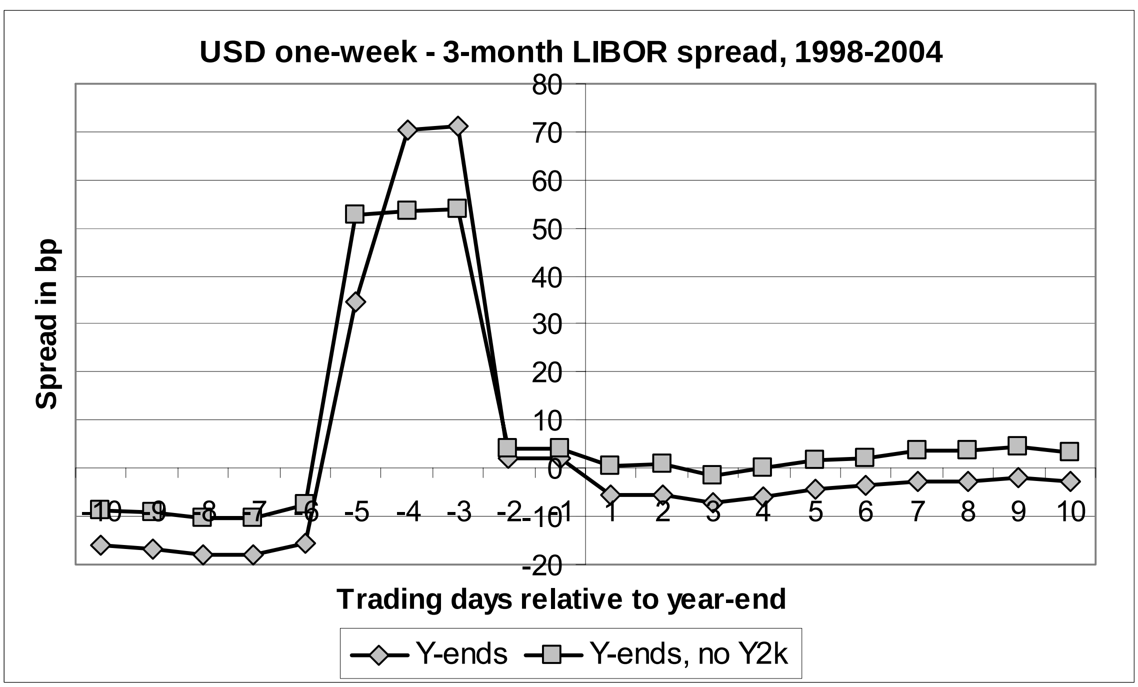

Figure 8. Average spread in basis points around year-ends between the one-week and three-month LIBOR for USD for year-ends 1998 through 2004.

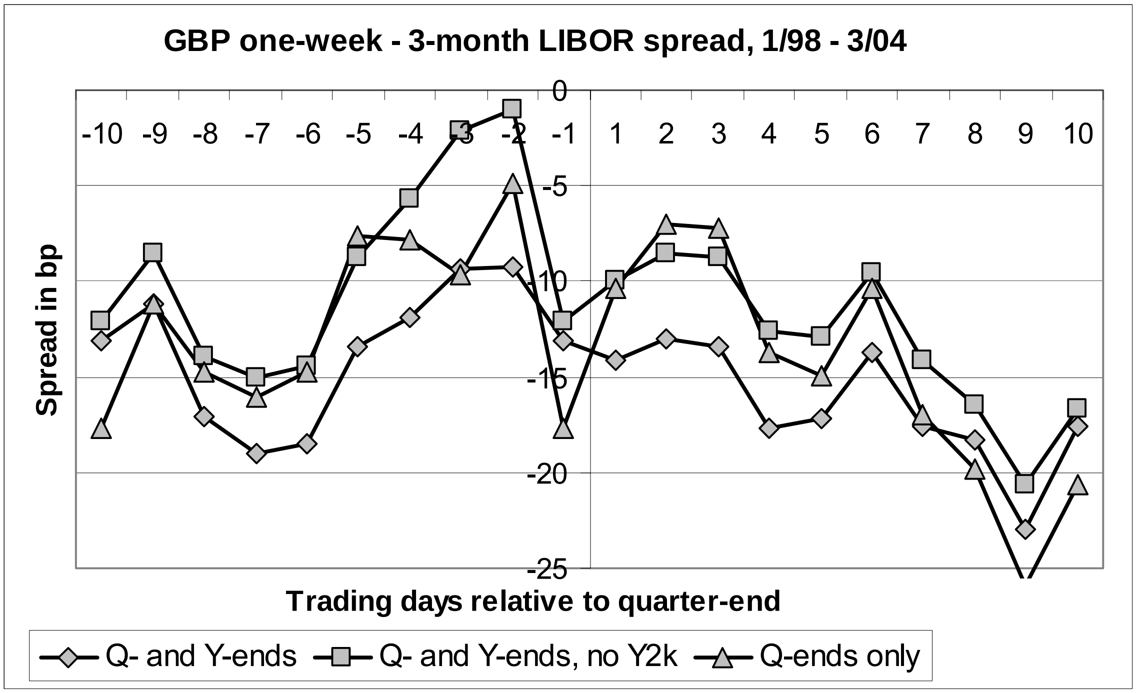

Figure 9. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for GBP over the period January 1998 through March 2004.

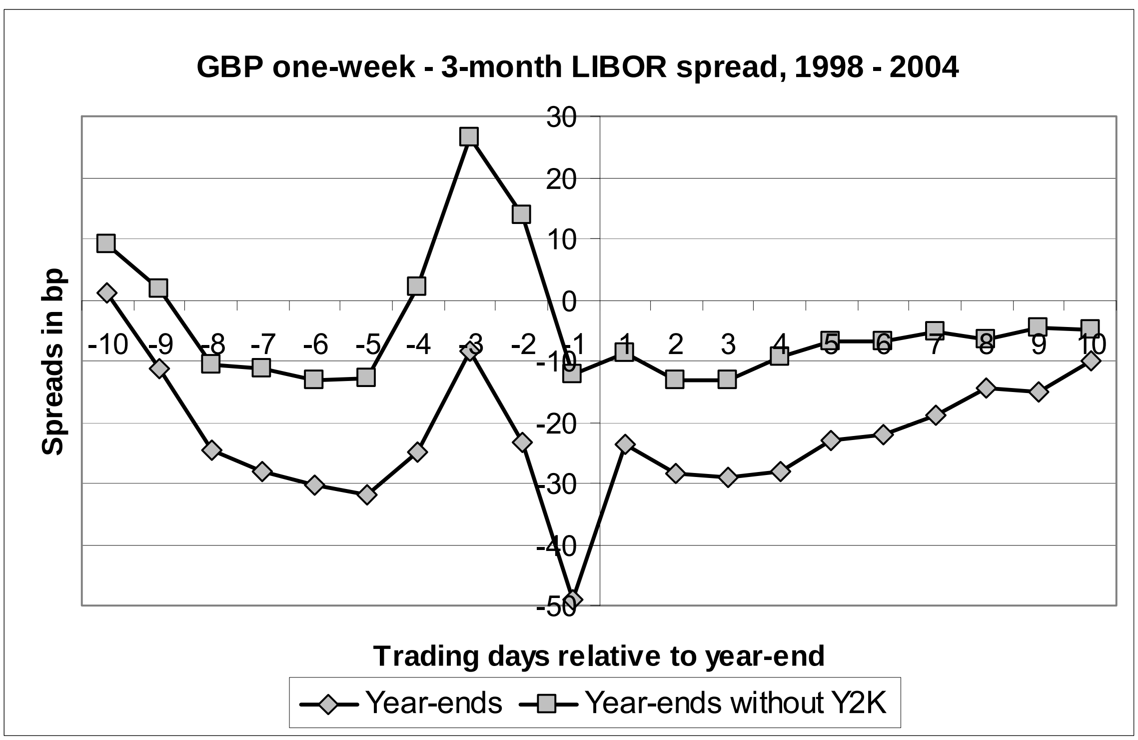

Figure 10. Average spread in basis points around year-ends between the one-week and three-month LIBOR for GBP for year-ends 1998 through 2004.

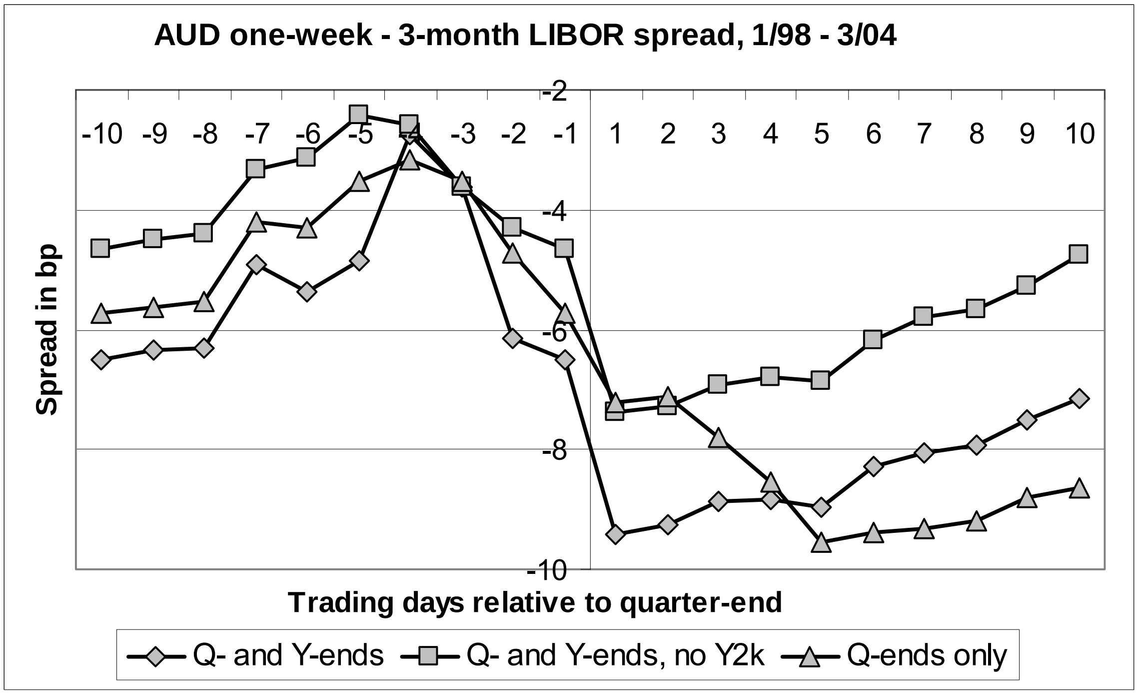

Figure 11. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for AUD over the period January 1998 through March 2004.

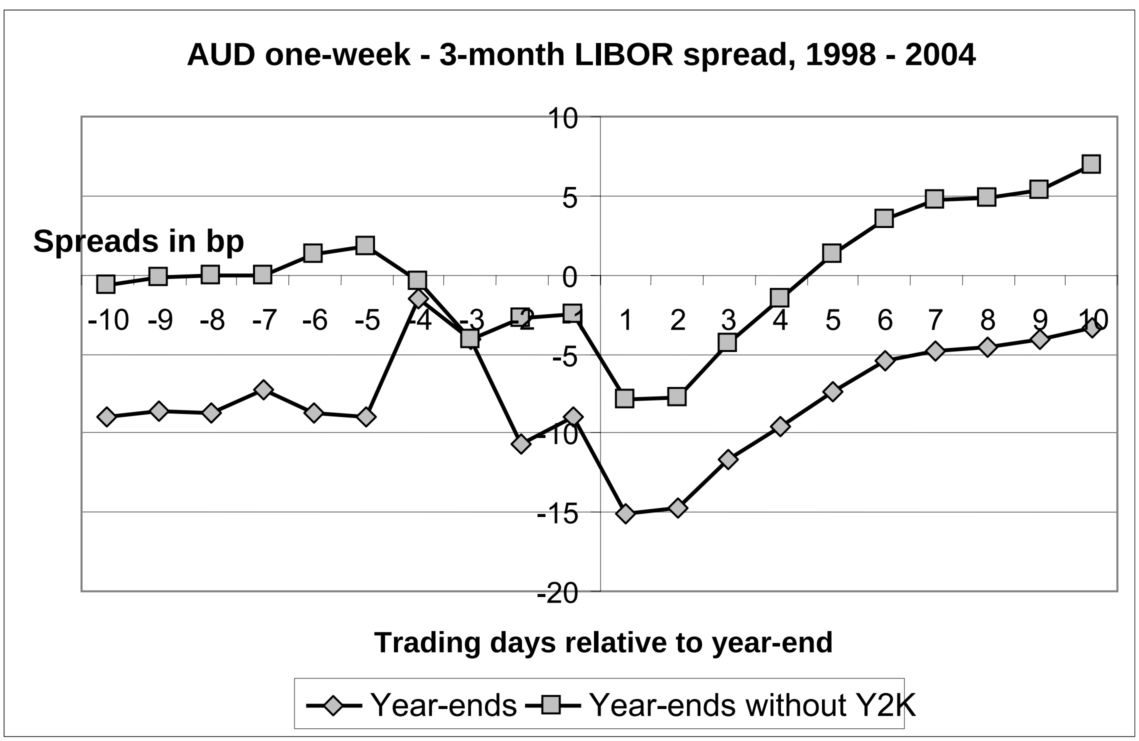

Figure 12. Average spread in basis points around year-ends between the one-week and three-month LIBOR for AUD for year-ends 1998 through 2004.

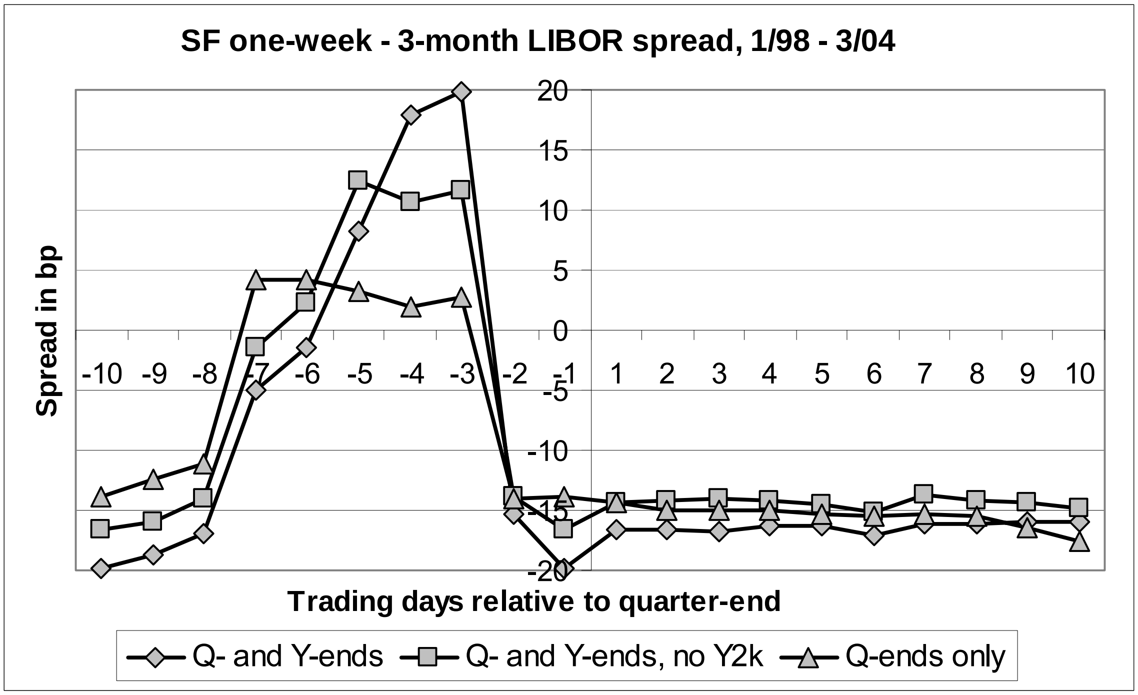

Figure 13. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for SF over the period January 1998 through March 2004.

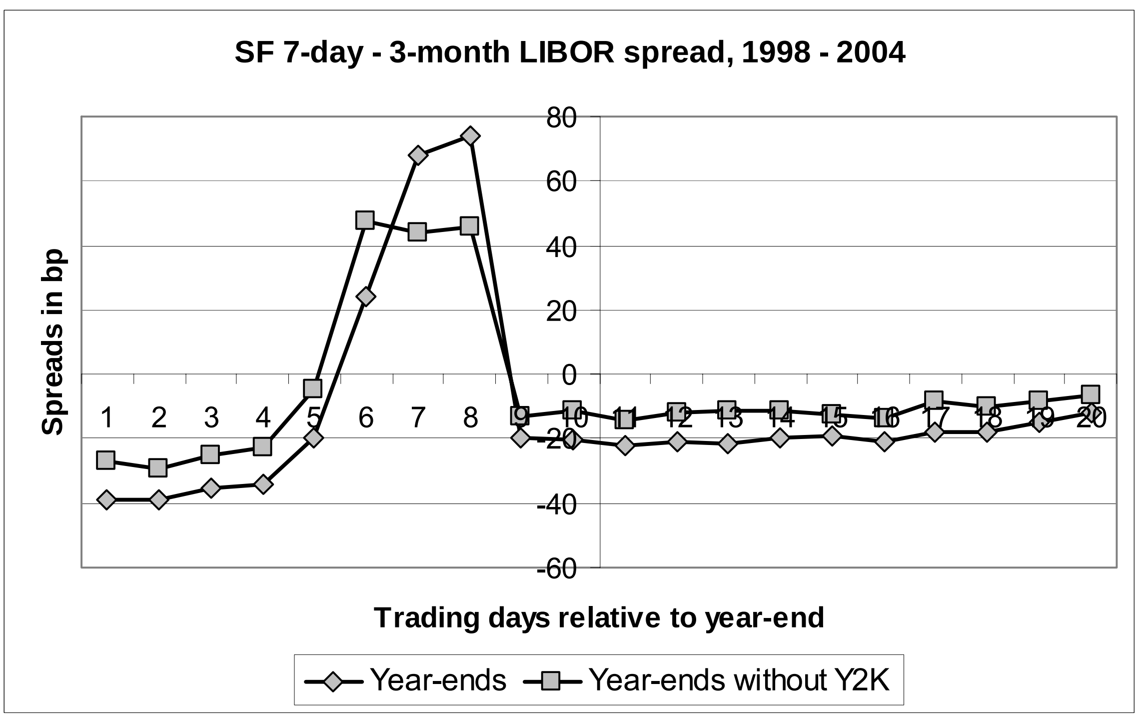

Figure 14. Average spread in basis points around year-ends between the one-week and three-month LIBOR for SF for year-ends 1998 through 2004.

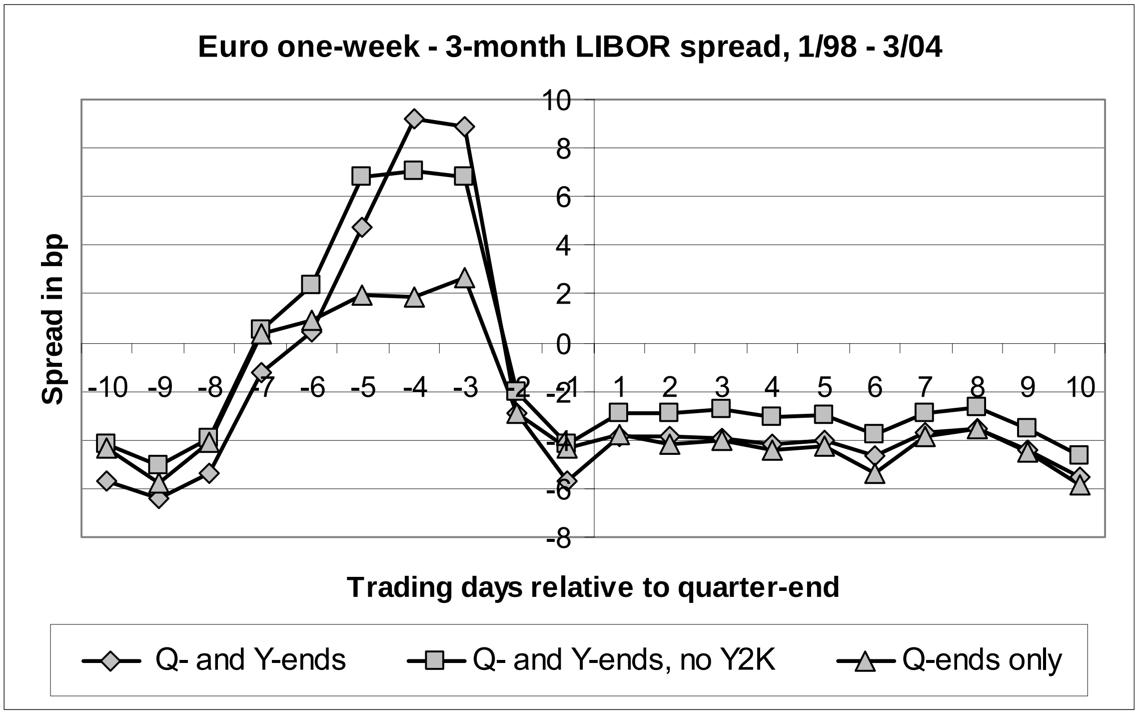

Figure 15. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for Euro over the period January 1998 through March 2004.

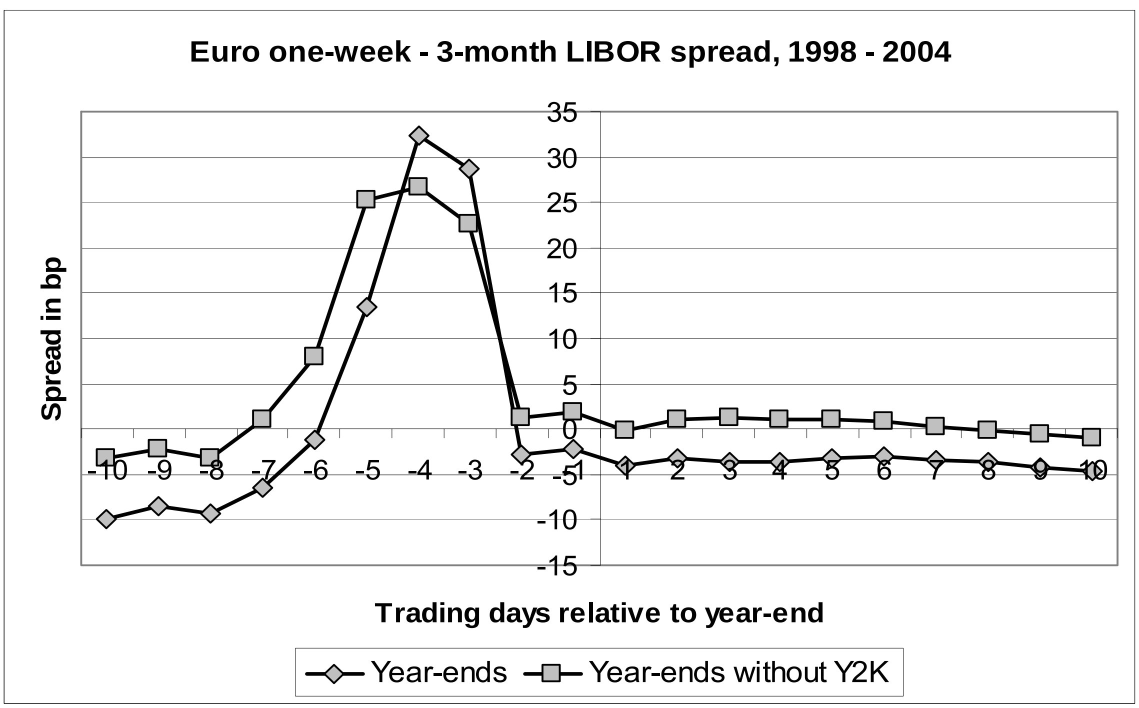

Figure 16. Average spread in basis points around year-ends between the one-week and three-month LIBOR for Euro for year-ends 1998 through 2004.

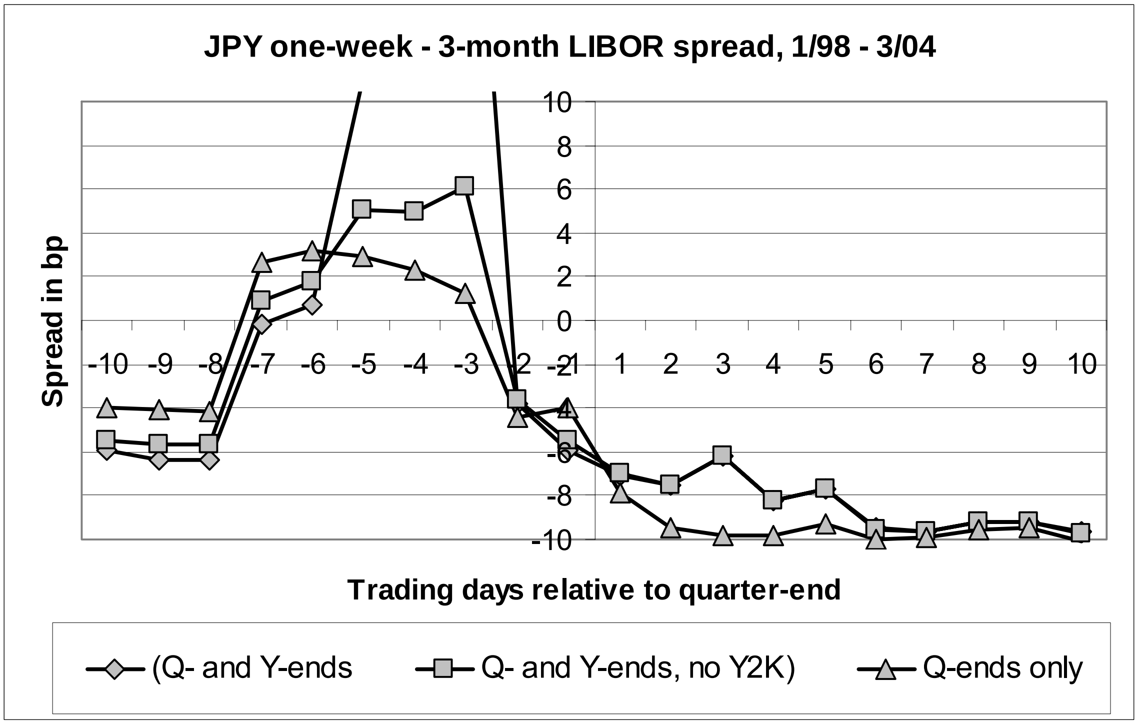

Figure 17. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for JPY over the period January 1998 through March 2004.

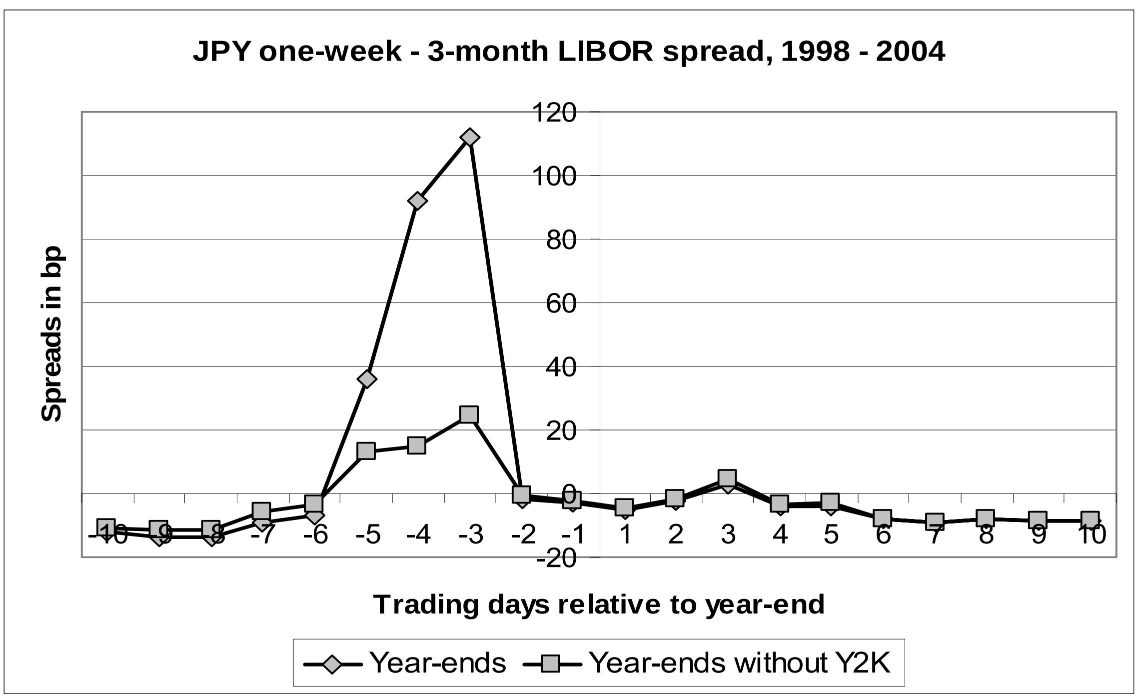

Figure 18. Average spread in basis points around year-ends between the one-week and three-month LIBOR for JPY for year-ends 1998 through 2004.

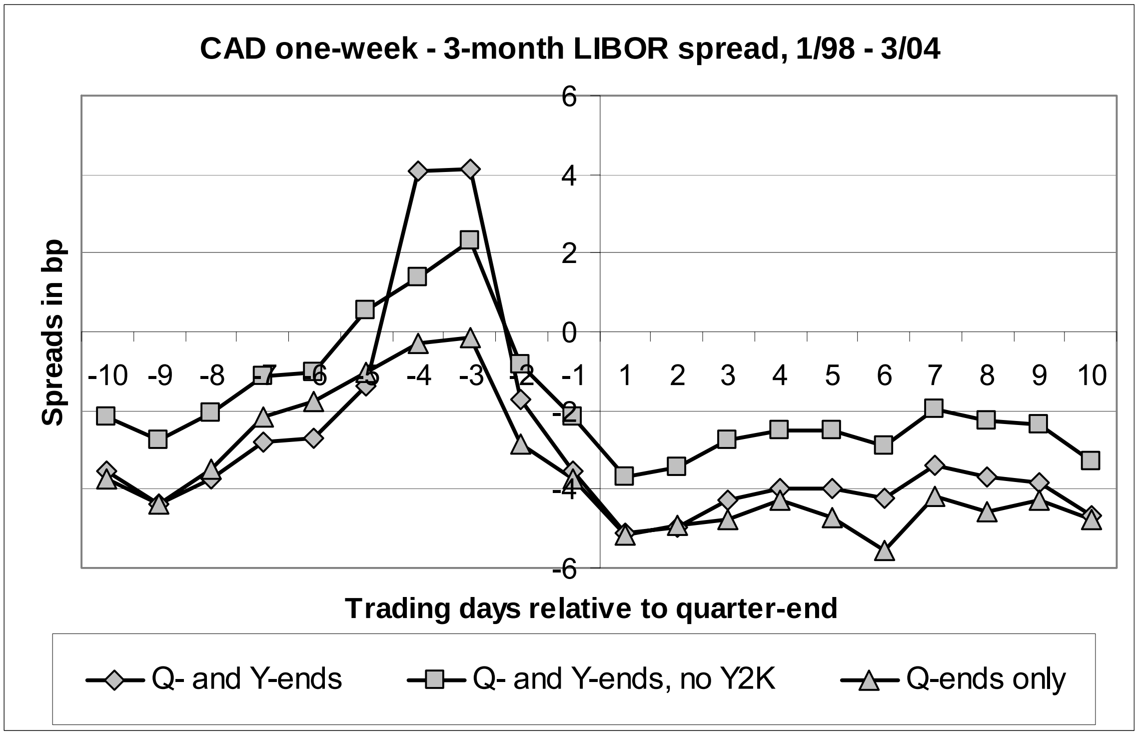

Figure 19. Average spread in basis points around quarter-ends between the one-week and three-month LIBOR for CAD over the period January 1998 through March 2004.

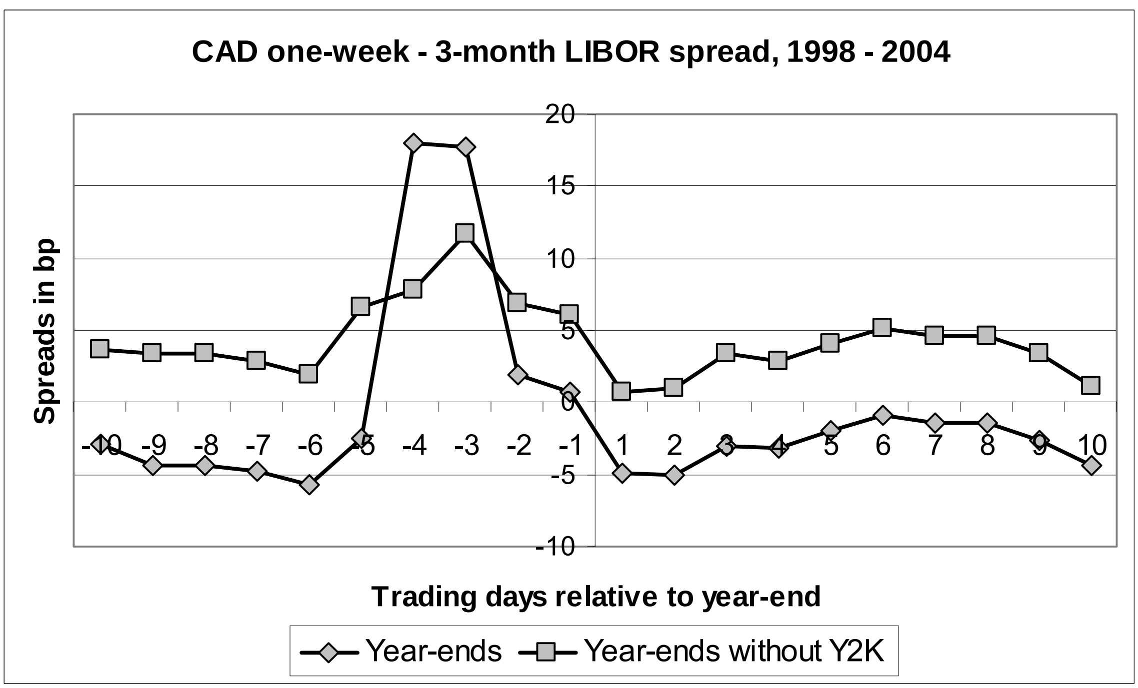

Figure 20. Average spread in basis points around year-ends between the one-week and three-month LIBOR for CAD for year-ends 1998 through 2004.

The patterns of spread changes for all seven currencies are consistent with the preferred habitat hypothesis; they are similar to that graphed for one-week repo spread in Figure 1 in Griffiths and Winters (1997).

Table 4 presents the output for regressions with the change in one-week LIBOR as a dependent variable. Again, the seven currencies for which the analysis was conducted are USD, GBP, AUD, Euro, SF, JPY, and CAD. Although one year of one-week LIBOR fixings is available for the Euro-in zone currencies, these data were not included in the analysis due to the small number of quarter-ends.

In Tables 4 and 5, statistically significant coefficients (at the 5% level or better) are in bold, and marginally significant coefficients (between the 5% and 10% level) are bold and italicized. Panel A of Table 4 contains the output from the model estimation over the entire data availability period, January 1998 through March 2004, without controlling for the Y2K. Panel B shows the results for the same time period with controls for the Y2K. We control for Y2K as a unique event in the sense that demand for liquidity was highest around it, which might have prompted the one-week rates to skyrocket in December 1999. A closer look at the one-week series supports this conjecture for six out of seven currencies. For example, one-week JPY LIBOR was 0.0825% on December 21, 1.875% on December 22, 5.25% on December 23, and 6.0625% on December 24, 1999. It fell to 0.1275% on December 29, the next business day. All other currencies with the exception of GBP also exhibited large, although not as striking, jumps in one-week LIBOR on these dates.

To isolate the effect of the Y2K, a separate set of turn-of-the-year dummy variables similar to those in equation (1) has been created. The coefficients of those dummies are not reported in the table not to stray the reader's attention off of "regular" year-ends.

Overall, the output supports preferred habitat for liquidity around quarter-ends. In the discussion of the results, we will focus more on the output presented in Panel B. The results are consistent with preferred habitat for the major world currencies: USD, Euro, SF, and JPY. For example, the coefficient of BQCR for USD (Panel B of Table 4) is 0.013. This means that the average change in the relative spread between one-week and three-month USD LIBOR is 0.013 higher on the two trading days preceding the day on which one-week LIBOR begins to mature in the next quarter than on "regular" days. If we assume that the three-month USD LIBOR stays unchanged at 3.930% (the mean of the three-month USD LIBOR over the period of analysis), the coefficient value of 0.013 would translate into a 5.1 basis points increase in the one-week rate *on each of these two days*.[9] The USD relative spread change is 0.018 lower on the last two days of the quarter than on other days; under the same assumptions, it is an equivalent of a 7.1 basis points decline in the one-week rate on each of these two days. Both coefficients are significant at the 1% level. The patterns of turn-of-the-quarter one-week – three-month LIBOR spread changes for Euro, SF, and JPY are similar. The daily basis point equivalent changes (again, assuming the average level of the three-month LIBOR) in one-week rates on each of the days corresponding respectively to BQCR and BQEND are: 3.5 and -3.9 basis points for Euro, 5.3 and -4.3 basis points for JPY, and 10.9 and -10.3 basis points in the case of SF. This pattern of spread changes is consistent with preferred habitat for liquidity and is not consistent with either variant of window dressing (see Table 3 on p.31). This result does not necessarily deny window dressing but it at least suggests that quarter-end preference for liquidity has more influence on the behavior of short-term interest rates for major world currencies around quarter-ends and year-ends.

[9] After solving for ΔR_t in $\Delta R_t/3.930 = 0.013$.

The British Pound and Canadian Dollar do not exhibit statistically significant evidence of quarter-end preferred habitat, although the signs of the coefficients are consistent with preferred habitat. In the case of the Australian Dollar, the decrease in the spread occurs before the quarter-end, while the increase is not fully captured by the dummy variables. A plot of AUD spread changes around quarter-ends in Figure 11 suggests that the increase in the spread may be too gradual to be captured by the dummy variables.

Turn-of-the year dummy variables' coefficients are much larger in magnitude than their turn-of-the quarter counterparts. This is not surprising since the highest need for liquidity normally arises before year-ends (see Ogden (1987)). Given a small number of observations (five year-ends after controlling for Y2K), the results are still statistically significant at the 5% level for the three currencies: USD, Euro, and SF. Although the year-end results in JPY and CAD are only marginally significant, the interest rate changes are large in magnitude and may be considered economically significant: Stigum (1990) notes that money market traders consider 10 to 20 basis points as an attractive arbitrage opportunity. The Pound Sterling shows a large significant decline in the spread prior to the year-end, while the preceding increase in the spread is comparable in magnitude to that decrease but is not statistically significant. It might be too gradual to be captured by the dummies. The Australian Dollar exhibits some statistically significant spread changes around the end of the year, but these changes are not consistent with any of the hypotheses.

The basis point equivalents of the relative spread changes prior to the year-end (again, under the assumption that the three-month LIBOR is at its average level and stays unchanged) are: for USD – an increase of 40 basis points (bp) on each of the two days covered by the BYCR dummy, followed by a 24-basis point decrease on each day covered by

the BYEND dummy; for Euro – a 14 bp increase and -12.3 bp decrease, for JPY –a 6.8 bp jump and a 7.2 bp fall, and for SF – a 47 bp spike and a 32 bp decrease, respectively. Overall, the fact that the participants in the London interbank market do not arbitrage away such profitable opportunities suggests that something is going on in the money market that prevents them from doing so. In the context of this study, the most likely explanation for this empirical phenomenon is that the high demand for liquidity around year-ends is not matched by the equally high supply, which drives up yields of short-term securities that mature in the new year.

Table 4. One-week LIBOR regressions results.

Panel A. Estimation output without controlling for Y2K

	USD		GBP		AUD		EURO		SF		JPY		CAD	
Variable	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**
Intercept	<0.001	0.743	<0.001	0.837	<0.001	0.575	<0.001	0.588	0.001	0.530	0.007	0.133	<0.001	0.765
LagDV	**–0.196**	**0.035**	**–0.171**	**0.000**	–0.050	0.281	**–0.204**	**0.002**	**–0.248**	**0.002**	0.022	0.897	–0.058	0.463
BQCR	**0.012**	**0.003**	–0.002	0.740	0.001	0.309	**0.008**	**0.000**	**0.059**	**0.000**	***0.122***	***0.099***	0.002	0.397
AQCR	0.002	0.147	0.008	0.238	0.001	0.279	***0.002***	***0.062***	**–0.012**	**0.006**	–0.046	0.204	**0.003**	**0.008**
BQEND	**–0.018**	**0.000**	0.001	0.910	**–0.004**	**0.011**	**–0.010**	**0.003**	**–0.057**	**0.000**	**–0.111**	**0.001**	**–0.005**	**0.014**
AQEND	***–0.002***	***0.093***	0.000	0.934	-0.001	0.416	-0.001	0.340	–0.006	0.182	**–0.023**	**0.004**	0.001	0.214
BYCR	**0.117**	**0.006**	0.005	0.793	0.008	0.340	**0.060**	**0.018**	**0.357**	**0.022**	0.548	0.117	0.024	0.194
AYCR	***–0.041***	***0.056***	0.037	0.146	***–0.008***	***0.081***	***–0.036***	***0.059***	**–0.183**	**0.035**	–0.649	0.264	–0.005	0.575
BYEND	**–0.065**	**0.014**	**–0.056**	**0.006**	0.000	0.971	**–0.037**	**0.029**	**–0.219**	**0.032**	–0.611	0.303	***–0.015***	***0.068***
AYEND	**–0.010**	**0.024**	0.011	0.654	**–0.007**	**0.010**	–0.001	0.586	0.010	0.346	0.009	0.787	0.000	0.859
F-statistic	72.0		7.9		3.6		44.0		63.1		22.5		11.3	
Adj. R^2	0.288		0.038		0.03		0.197		0.262		0.109		0.056	

Panel B. Y2K controls included

	USD		GBP		AUD		EURO		SF		JPY		CAD	
Intercept	<0.001	0.823	<0.001	0.824	<0.001	0.595	<0.001	0.735	<0.001	0.750	0.004	0.314	<0.001	0.856
LagDV	**–0.240**	**0.014**	**–0.187**	**0.000**	***–0.119***	***0.052***	**–0.349**	**0.001**	**–0.364**	**0.000**	**-0.766**	**0.000**	***–0.225***	***0.073***
BQCR	**0.013**	**0.002**	–0.002	0.725	0.001	0.274	**0.010**	**0.000**	**0.065**	**0.000**	**0.238**	**0.003**	0.002	0.262
AQCR	0.002	0.126	0.008	0.227	0.001	0.251	**0.002**	**0.040**	**–0.012**	**0.001**	–0.059	0.276	**0.003**	**0.008**
BQEND	**–0.018**	**0.000**	0.001	0.906	**–0.004**	**0.009**	**–0.011**	**0.001**	**–0.061**	**0.000**	**–0.192**	**0.000**	**–0.005**	**0.009**
AQEND	–0.002	0.089	–0.001	0.929	-0.001	0.333	-0.001	0.384	–0.006	0.225	**–0.030**	**0.001**	0.001	0.357
BYCR	**0.102**	**0.012**	0.017	0.439	-0.001	0.631	**0.040**	**0.000**	**0.278**	**0.040**	***0.303***	***0.081***	**0.010**	**0.024**
AYCR	–0.032	0.140	0.041	0.117	**–0.005**	**0.032**	–0.023	0.156	–0.115	0.135	–0.066	0.765	0.004	0.480
BYEND	**–0.061**	**0.020**	**–0.053**	**0.009**	**0.004**	**0.037**	**–0.035**	**0.007**	**–0.190**	**0.029**	***–0.325***	***0.103***	***–0.007***	***0.076***
AYEND	**–0.005**	**0.038**	–0.009	0.698	**–0.006**	**0.006**	0.001	0.526	**0.026**	**0.000**	0.010	0.776	0.003	0.125
F-statistic	63.9		8.5		16.1		66.2		84.4		207.2		31.3	
Adj. R^2	0.324		0.054		0.102		0.332		0.388		0.611		0.188	

Turn-of-the-quarter and turn-of-the-year effect in LIBOR for major world currencies in the one-week maturity. The parameters are estimated using Equation (1):

$$\Delta Rsp_t = \alpha_0 + \alpha_1(\Delta Rsp_{t-1}) + \alpha_2 BQCR + \alpha_3 AQCR + \alpha_4 BQEND + \alpha_5 AQEND + \alpha_6 BQCR + \alpha_7 AQCR + \alpha_8 BQEND + \alpha_9 AQEND + \varepsilon_t$$

The dependent variable is the daily change in the relative spread between the one-week and three-month LIBOR for a given currency; the dummy variables BQCR, AQCR, BQEND, and AQEND are designed to capture changes in the relative spread on days surrounding ends of quarters one, two, and three, while the dummy variables BYCR, AYCR, BYEND, and AYEND perform the same function around year-ends.

1.3.2 One-month Results

Figures 21 and 22 present the dynamics of the spread between the one-month and three-month USD LIBOR around quarter-ends and year-ends over the period of the analysis (1/87 through 3/04). It is clear that the spread increases on the days before the maturity of one-month LIBOR starts to span the end of the quarter; this increase is sustained until the second to last trading day of the quarter. Figure 21 is very similar to Figure 2 in Griffiths and Winters (2004).

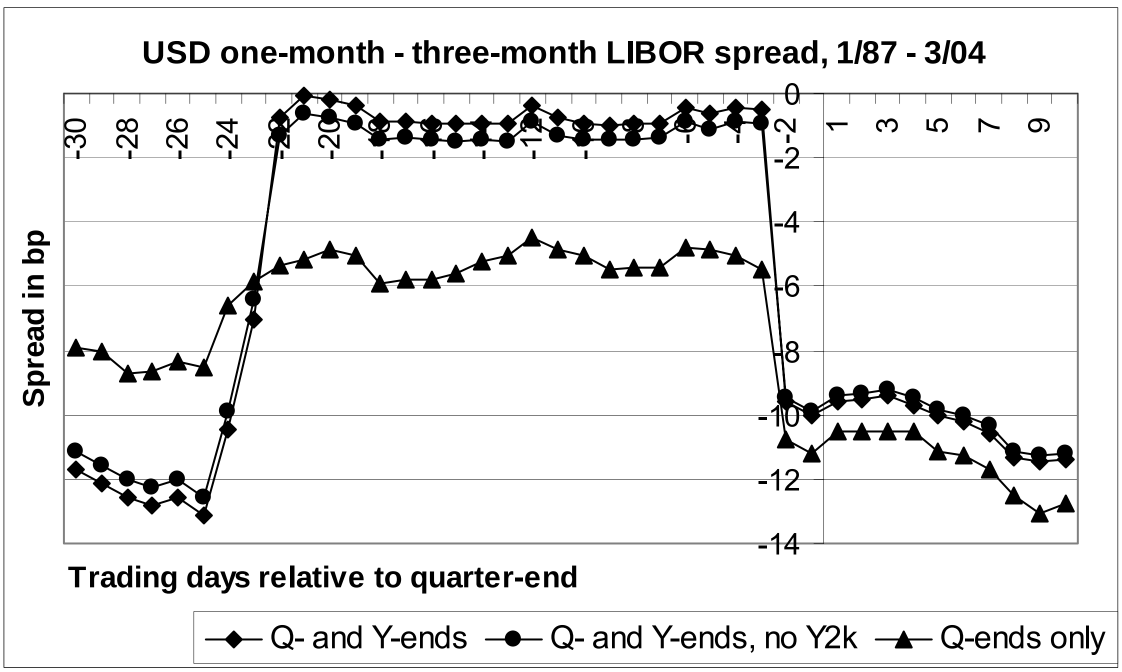

Figure 21. Average spreads between the one-month and three-month U.S. Dollar LIBOR around quarter- and year-ends, over the period January 1987 through March 2004.

It is quite evident from Figure 21 that the year-end effect in one-month USD LIBOR is much larger in magnitude than the quarter-end effect. Figure 22 isolates year-ends from quarter-ends. Along with Figure 21, it emphasizes the magnitude of the turn-of-the year effect which by far exceeds that of the regular turn-of-the-quarter effect in one-month USD LIBOR. It also demonstrates that the turn of the Year 2000 was not a major factor

contributing into the dynamics of the spread behavior around year-ends. This is opposite of what we have observed for one-week LIBOR. The minimal Y2K influence on this data series may be due to the longer maturity of the instrument reflecting the lack of urgency to borrow before Y2K. Also, one-month spreads have been averaged over 17 years of data available for this maturity. Any particular observation has to truly stand out to significantly affect the average. This is apparently not the case with Y2K for one-month USD LIBOR; it is also true for the other six currencies with the data available around Y2K. Therefore, only one graph is reported for each of the other six currencies. It contains two series: one depicting spread changes around "regular" quarter-ends and the other plotting spreads around year-ends. Year-end effects are much larger in magnitude than quarter-end effects.

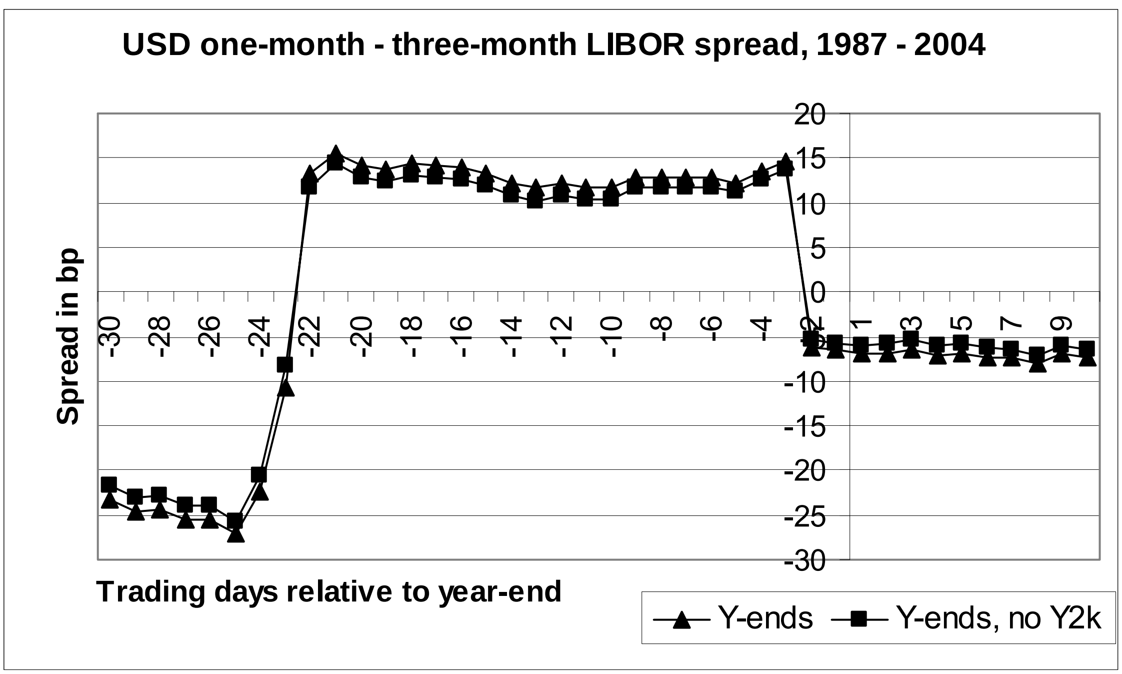

Figure 22. Average spreads between the one-month and three-month U.S. Dollar LIBOR around year-ends.

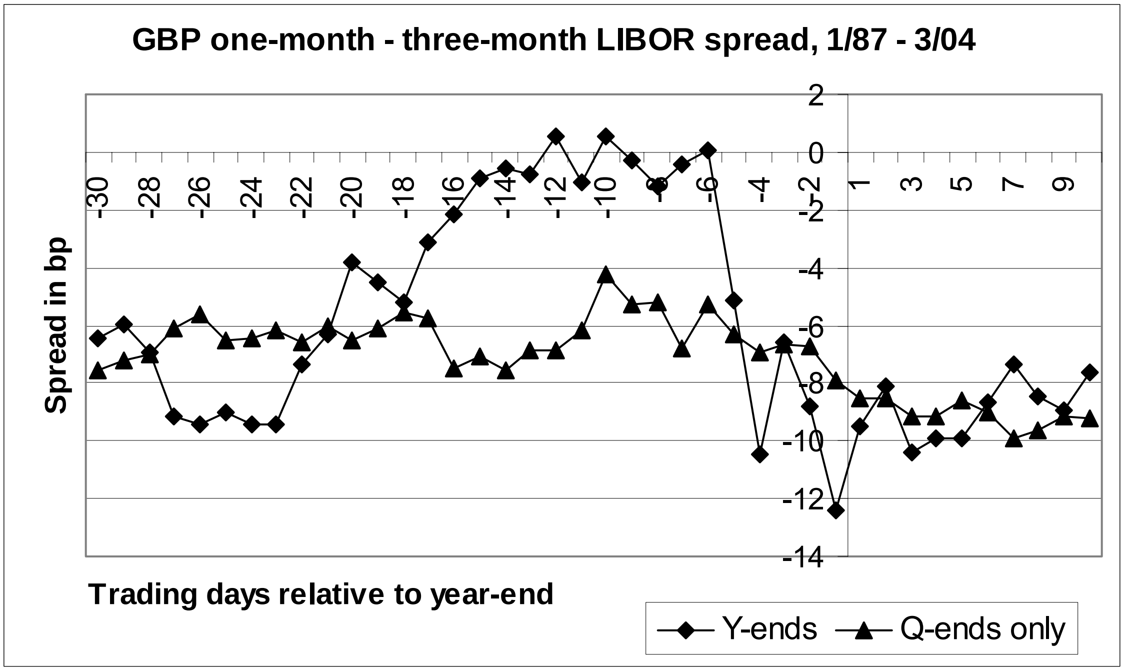

Figure 23. Average spreads between the one-month and three-month Pound Sterling LIBOR around quarter- and year-ends, over the period January 1987 through March 2004.

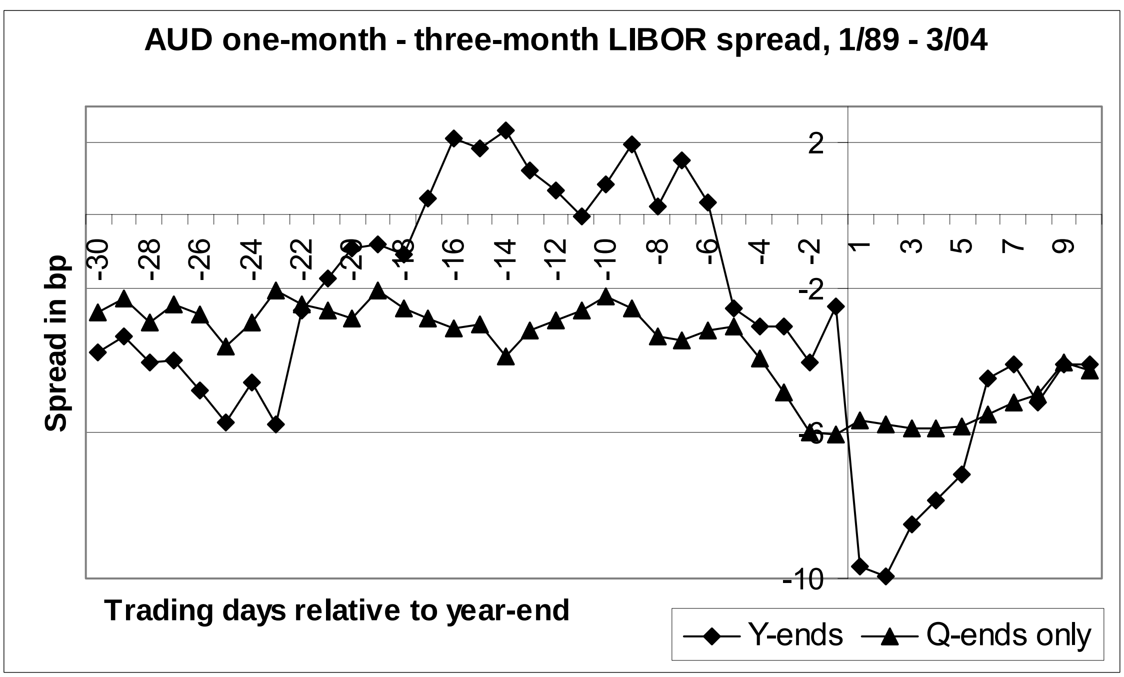

Figure 24. Average spreads between one-month and three-month Australian Dollar LIBOR around quarter- and year-ends, over the period January 1989 through March 2004.

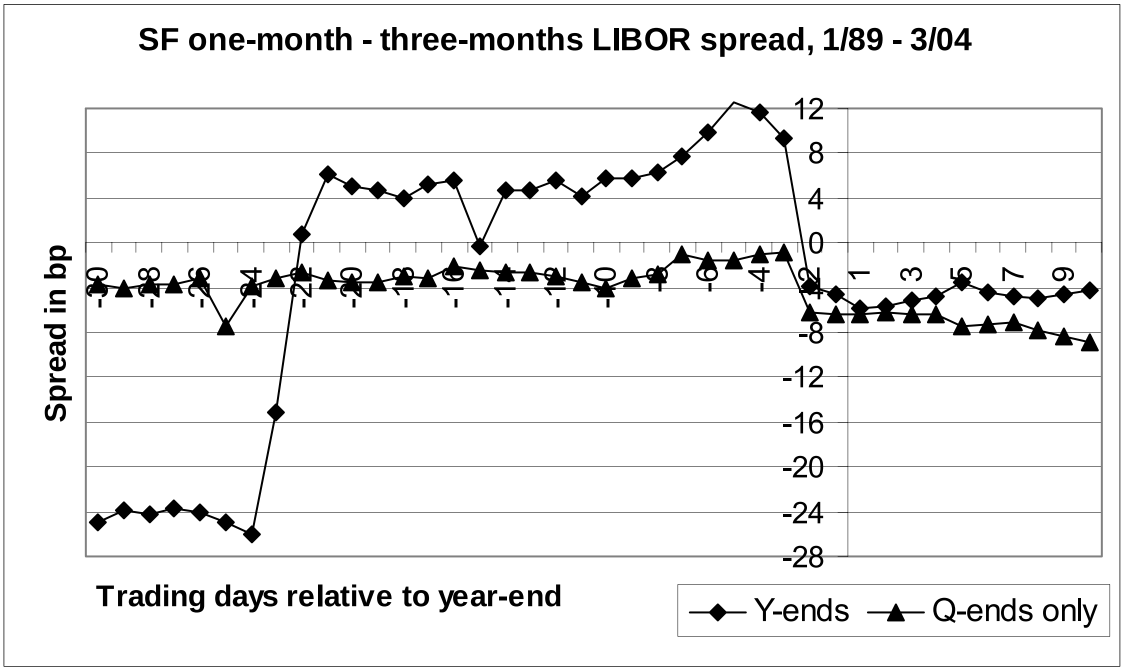

Figure 25. Average spreads between the one-month and three-month Swiss Frank LIBOR around quarter- and year-ends, over the period January 1989 through March 2004.

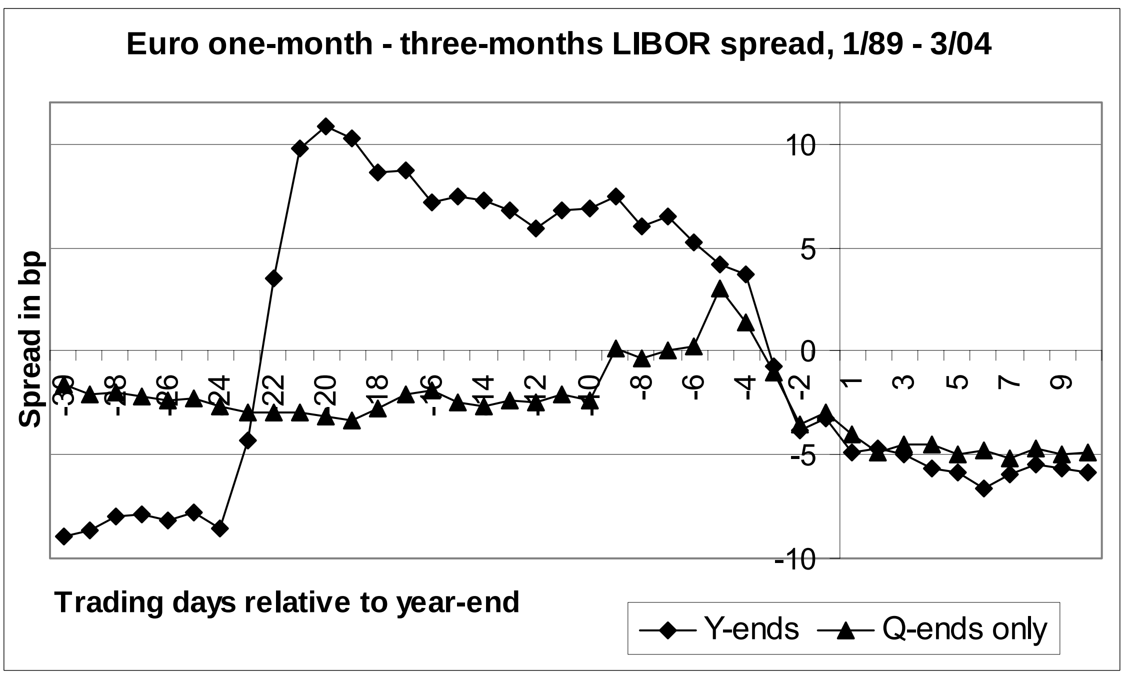

Figure 26. Average spreads between the one-month and three-month Euro LIBOR around quarter- and year-ends, over the period January 1989 through March 2004.

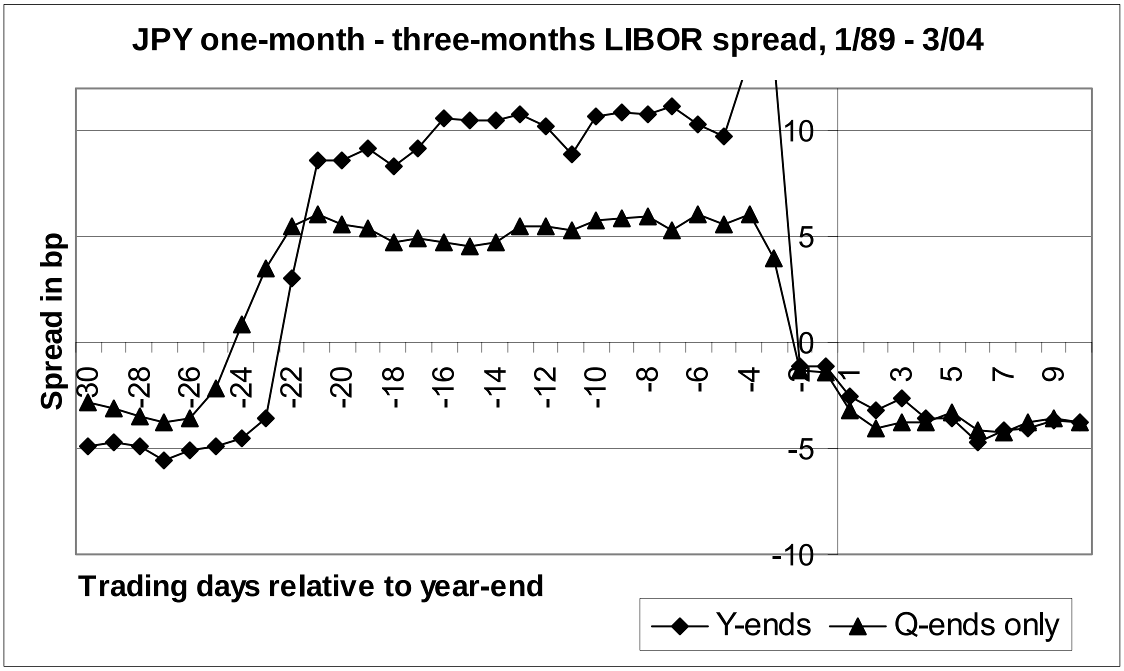

Figure 27. Average spreads between the one-month and three-month JPY LIBOR around quarter- and year-ends, over the period January 1989 through March 2004.

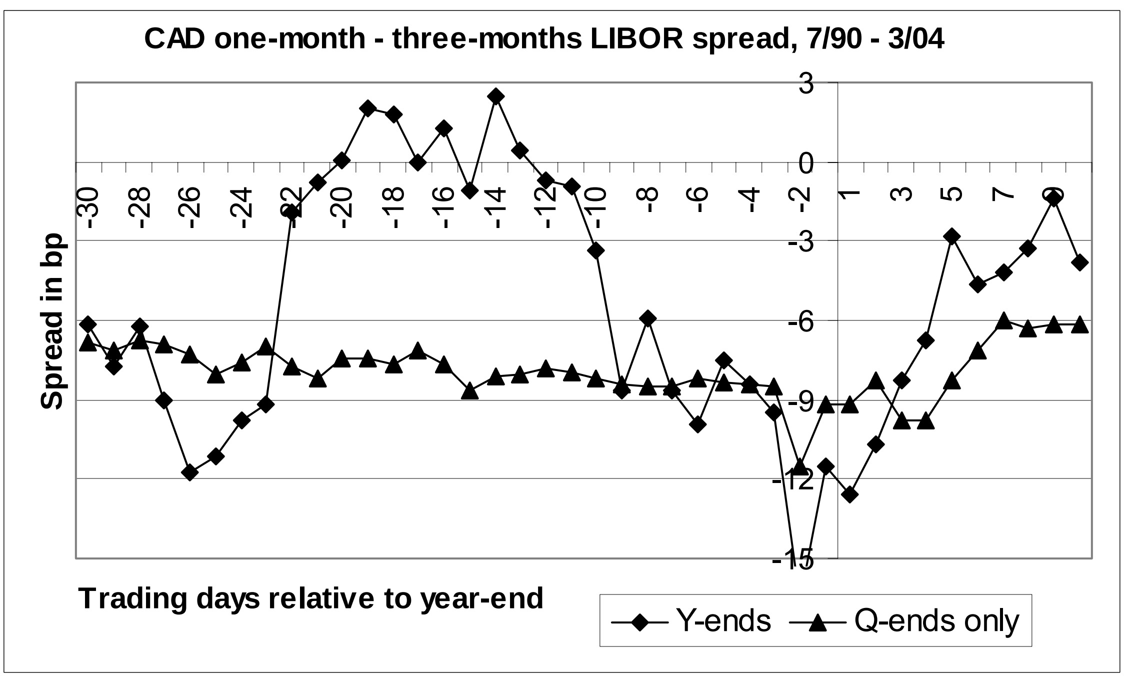

Figure 28. Average spreads between the one-month and three-month Canadian Dollar LIBOR around quarter- and year-ends, over the period July 1990 through March 2004.

Figures 21 through 28 demonstrate that the year-end increase in spreads exists in all seven currencies. While these increases start on day -22 or -23 relative to the year-end (which typically correspond to the BYCR dummy) in all currencies, the subsequent decrease is less uniform across the currencies, although it always starts prior to the calendar year-end. For example, in the case of the Canadian Dollar, a sharp decrease occurs on day -9, while for the Pound Sterling and the Australian Dollar the decline starts on day -5. For the rest of the currencies, the larger spreads are sustained through the third-to-last or second-to-last day of the year.[10]

Table 5 contains the regression output for one-month LIBOR for all 11 currencies included in the analysis. Panels A and B have been constructed similarly to those in Table 4: they present the results for the same seven currencies with and without the dummies controlling for Y2K. Panel C contains the regression output for the four Euro-in zone currencies. LIBOR fixings for these currencies were ceased at the end of 1998, which renders the Y2K problem irrelevant for them.

The output presented in Table 5 provides support for preferred habitat at both quarter-ends and year-ends. The magnitude of the year-end effect differs across currencies: the increase and the subsequent decrease are higher in USD, Euro, JPY, and SF than in other currencies, although in the case of the Euro the decrease is rather gradual. The turn-of-the-year effect consistent with preferred habitat for liquidity is statistically significant but small in GBP. In the case of CAD, the decrease may not have been captured by the dummy variables: as evident from Figure 28, a drop in the spread occurs in CAD prior to the days covered by the dummy BYEND. Among the currencies replaced by the Euro, DM and ITL

[10] The two-day spike in the spread for JPY on day -4 is due to Y2K; this is its only noticeable influence for the one-month LIBOR across all currencies.

exhibit the pattern of one-month interest rate changes consistent with the turn-of-the-year preferred habitat for liquidity. None of the currencies exhibit behavior consistent with either idle cash or flight-from-risk window dressing.

One easily noticeable feature of the output reported in Table 5 is that the magnitude of the quarter-end effect in the one-month LIBOR is much smaller than in the one-week LIBOR, while the year-end effect is large and statistically significant in both one-week and one-month maturities. In the one-month maturity, only JPY demonstrates a turn-of-the quarter effect that is both economically and statistically significant. The turn-of-the-quarter coefficients for USD are statistically significant but small. For Euro and SF, the turn-of-the quarter coefficients lack statistical significance. These differences in quarter-end results between one-week and one-month maturities suggest that banks and their customers start preparing for the end of the year well in advance (and this is reflected in the turn-of-the-year effect in the one-month interest rate), while the preparation to meet quarter-end obligations starts relatively close to the calendar quarter-end. Therefore, we observe a strong turn-of-the-year effect in both one-week and one-month LIBOR, but a strong turn-of-the-quarter effect only in the one-week maturity.

The results reported here are in line with the findings of Griffiths and Winters (2004), although the methods are different. This may imply robustness of the results to different methods.[11]

[11] The results do not change qualitatively when other methods are used.

Table 5. One-month LIBOR results.

Panel A. Estimation output without controlling for Y2K

	USD		GBP		AUD		EURO		SF		JPY		CAD	
Variable	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**
Intercept	***<0.001***	***0.060***	<0.001	0.984	<0.001	0.953	<0.001	0.940	<0.001	0.716	<0.001	0.635	<0.001	0.626
LagDV	**-0.293**	**0.000**	**-0.126**	**0.008**	**-0.147**	**0.000**	**-0.173**	**0.006**	**-0.270**	**0.007**	-0.033	0.640	**-0.124**	**0.028**
BQCR	**0.004**	**0.000**	0.000	0.563	**0.001**	**0.042**	0.000	0.955	0.000	0.963	**0.042**	**0.000**	0.001	0.131
AQCR	0.000	0.678	0.000	0.634	0.000	0.805	***-0.001***	***0.056***	-0.001	0.314	0.000	0.869	-0.001	0.177
BQEND	**-0.006**	**0.000**	-0.001	0.222	**-0.001**	**0.048**	**-0.003**	**0.046**	**-0.015**	**0.000**	**-0.032**	**0.000**	-0.001	0.624
AQEND	**0.001**	**0.011**	-0.001	0.283	0.000	0.992	-0.001	0.291	0.001	0.522	***0.004***	***0.063***	0.001	0.238
BYCR	**0.048**	**0.000**	**0.002**	**0.029**	***0.004***	***0.089***	**0.016**	**0.000**	**0.072**	**0.000**	***0.102***	***0.097***	***0.010***	***0.057***
AYCR	-0.001	0.268	0.003	0.210	0.001	0.639	0.001	0.387	-0.003	0.168	0.001	0.834	0.001	0.601
BYEND	**-0.022**	**0.000**	**-0.004**	**0.024**	-0.001	0.604	-0.003	0.266	**-0.037**	**0.000**	-0.101	0.117	-0.002	0.381
AYEND	***-0.002***	***0.060***	0.002	0.154	**-0.002**	**0.012**	0.000	0.961	-0.002	0.451	0.002	0.727	0.001	0.501
F-statistic	157.0		10.2		12.7		29.5		76.7		27.6		9.3	
Adj. R^2	0.244		0.019		0.027		0.062		0.150		0.058		0.021	

Panel B. Y2K controls are included

	USD		GBP		AUD		EURO		SF		JPY		CAD	
Intercept	***<0.001***	***0.057***	<0.001	0.987	<0.001	0.942	<0.001	0.928	<0.001	0.696	<0.001	0.705	<0.001	0.619
LagDV	**-0.309**	**0.000**	**-0.138**	**0.004**	**-0.177**	**0.000**	**-0.201**	**0.000**	**-0.353**	**0.001**	**-0.494**	**0.000**	**-0.160**	**0.011**
BQCR	**0.004**	**0.000**	<0.001	0.568	**0.001**	**0.041**	<0.001	0.947	<0.001	0.947	**0.057**	**0.000**	0.001	0.137
AQCR	<0.001	0.699	<0.001	0.650	<0.001	0.811	***-0.001***	***0.053***	-0.001	0.294	0.004	0.189	-0.001	0.191
BQEND	**-0.007**	**0.000**	-0.001	0.219	**-0.001**	**0.041**	**-0.003**	**0.039**	**-0.015**	**0.000**	**-0.047**	**0.000**	-0.001	0.599
AQEND	**0.001**	**0.011**	-0.001	0.278	<0.001	0.988	-0.001	0.293	0.001	0.548	**0.006**	**0.011**	0.001	0.227
BYCR	**0.046**	**0.000**	**0.002**	**0.016**	***0.002***	***0.081***	**0.012**	**0.000**	**0.059**	**0.000**	**0.070**	**0.019**	**0.005**	**0.006**
AYCR	-0.001	0.295	0.001	0.521	0.001	0.533	0.002	0.412	-0.004	0.153	-0.000	0.914	0.003	0.106
BYEND	**-0.020**	**0.000**	**-0.004**	**0.042**	0.001	0.518	**-0.005**	**0.018**	**-0.031**	**0.000**	**-0.053**	**0.004**	-0.001	0.628
AYEND	-0.001	0.128	0.001	0.428	**-0.002**	**0.025**	<0.001	0.741	-0.000	0.861	0.003	0.659	**0.002**	**0.041**
F-statistic	119.6		11.4		21.0		42.3		99.4		242.1		17.9	
Adj. R^2	0.261		0.030		0.063		0.122		0.249		0.449		0.060	

Panel C. Euro-in zone currencies								
	DM		**FF**		**SP**		**ITL**	
Variable	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**
Intercept	<0.001	0.207	<0.001	0.911	<0.001	0.927	<0.001	0.900
LagDV	**-0.352**	**0.000**	***-0.244***	***0.052***	-0.363	0.128	-0.127	0.365
BQCR	**0.002**	**0.001**	<0.001	0.678	0.001	0.698	<0.001	0.894
AQCR	<0.001	0.643	-0.001	0.168	-0.002	0.178	0.001	0.278
BQEND	-0.002	0.230	0.001	0.699	0.000	0.857	-0.001	0.414
AQEND	-0.001	0.367	-0.002	0.444	-0.001	0.333	0.001	0.568
BYCR	**0.030**	**0.000**	0.002	0.303	0.002	0.260	**0.011**	**0.011**
AYCR	-0.001	0.554	**0.005**	**0.037**	0.001	0.385	0.001	0.457
BYEND	**-0.009**	**0.000**	-0.001	0.506	***-0.002***	***0.078***	**-0.009**	**0.013**
AYEND	-0.001	0.251	0.004	0.265	0.001	0.383	<0.001	0.836
F-statistic	76.7		18.6		36.3		6.0	
Adj. R-squared	0.183		0.059		0.129		0.020	

Turn-of-the-quarter and turn-of-the-year effect in LIBOR for major world currencies in the one-month maturity. The parameters are estimated using Equation (1):

$$\Delta Rsp_t = \alpha_0 + \alpha_1(\Delta Rsp_{t-1}) + \alpha_2 BQCR + \alpha_3 AQCR + \alpha_4 BQEND + \alpha_5 AQEND + \alpha_6 BQCR + \alpha_7 AQCR + \alpha_8 BQEND + \alpha_9 AQEND + \varepsilon_t$$

The dependent variable is the daily change in the relative spread between the one-month and three-month LIBOR for a given currency; the dummy variables BQCR, AQCR, BQEND, and AQEND are designed to capture changes in the relative spread on days surrounding ends of quarters one, two, and three, while the dummy variables BYCR, AYCR, BYEND, and AYEND perform the same function around year-ends.

The results for the two-month LIBOR are similar to those for the one-month LIBOR, although they are significantly less pronounced. To save space, they are not reported. The three-month LIBOR and longer maturities do not exhibit the turn-of-the quarter effects in any of the currencies.

1.3.3 "Large" versus "Small" Currencies

As is evident from the empirical results reported above, four of the major world currencies – USD, Euro, JPY, and SF – have larger quarter-end and year-end effects in the one-week maturity than the currencies serving smaller markets – GBP, CAD, and AUD. The

same is true about the one-month maturity, where the same four currencies (USD, Euro, JPY, and SF) and also DM (which played the role of the European currency before the introduction of the Euro) exhibit much larger year-end effects than other currencies. The only “small-market” currency with the statistically significant year-end effect consistent with preferred habitat for liquidity is ITL.

The year-end effects in GBP and CAD are less significant, both statistically and economically, than in the major four currencies (five in the case of the one-month maturity), but they are still noticeable; there is no doubt about the presence of the year-end effect in these two currencies. The currency that exhibits the least consistency with preferred habitat for liquidity or window dressing is AUD. Two Euro-in zone currencies, FF and SP, also have year-end changes in yields that are close to nonexistent.

As noted above, currencies that exhibit statistically and economically significant quarter-end and year-end effects tend to be “large” in terms of the market size of the countries they serve. The U.S., the European Union, and Japan represent three of the four world’s largest economies (measured by GDP), as well as the world’s three largest importers and exporters. Switzerland’s economy is not comparable to these three in terms of size; however, total assets of Swiss banks exceed those of Australia and Canada combined and rank among highest in the world. This may explain the presence of the quarter-end and year-end preferred habitat for liquidity in the Swiss Franc.

The Pound Sterling exhibits the year-end effect lacking statistical significance. This is somewhat surprising in that the U.K. is one of the largest economies in the world in terms of GDP, export, and the bank assets. The graphical analysis of the data (Figures 9, 10, and 23) suggests that the regression may not detect a statistically significant effect due to the more

gradual changes in the spread than in other currencies: the relative spread tends to increase over four or five days instead of two. Also, GBP LIBOR has been noticeably more volatile over the study period than other interest rates (see Figures 4 through 6). High variance in the data makes confidence intervals wider and the null hypothesis of no quarter-end and no year-end effects less likely to be rejected.

The apparent division between big-market and smaller-market currencies in terms of the significance and magnitude of the quarter-end and year-ends preferred habitat effects suggests that the demand for the major world currencies – U.S. Dollar, Euro, Japanese Yen, Swiss Franc, and, to a lesser degree, Pound Sterling, – peaks prior to quarter-ends and especially year-ends. Banks' behavior (expressed in the London Interbank Offer Rate) is evidently consistent with the need of bank customers to meet their quarter-end and year-and cash obligations (business customers) or to make large expenditures (such as year-end shopping for individuals). While it is not exactly clear what group of bank customers contributes the most to these calendar effects, large multinational corporations involved in export and import operations, especially in the U.S., Europe, and Japan, must be one of the major forces.

1.4 Conclusion

The first part of this study extends the literature on calendar-time effects in the money markets by examining turn-of-the-quarter and turn-of-the-year effects in short-term interest rates for various currencies. This is done by using London Interbank Offer Rates for short-term loans in different currencies. The quarter-end and year-end effects behavior in very short-term rates (one-week and one-month maturities) is consistent with the preferred habitat for liquidity. These quarter-end and year-end effects are particularly large and statistically significant in the U.S dollar and three of the other major world currencies: the Euro, the Japanese Yen, and the Swiss Franc. The results for the Pound Sterling and the Canadian Dollar are less pronounced and less significant, but still consistent with the preferred habitat for liquidity. Among the currencies replaced by the Euro, the German Mark and the Italian Lira demonstrated changes in interest rates consistent with the year-end preferred habitat for liquidity in the one-month maturity.

The results suggest that the quarter-end and year-end demand for liquidity is driven by large multinational corporations operating with major world currencies (USD, Euro, JPY, SF, and, before the introduction of the Euro, DM). Behavior of short-term yields of the currencies that do not exhibit signs of preferred habitat (AUD, FF, and SP) is not consistent with the competing hypothesis, window dressing.

Overall, the findings further emphasize the importance of investors' preference for liquidity and its influence on money market yields. The evidence presented in this study does not refute window dressing by financial institutions; rather, it suggests that window dressing is not a dominating force in the money markets around quarter-ends and year-ends.

PART 2. TESTS OF THE EXPECTATIONS HYPOTHESIS IN THE PRESENCE OF PREFERRED HABITAT FOR LIQUIDITY.

2.1 Theories of the Term Structure of Interest Rates

The term structure of interest rates, or the relations among yields of securities with different time to maturity, has been of major interest to economists. One of the oldest models in finance is the expectations hypothesis introduced by Fisher (1896). The expectations hypothesis (EH) relates short-term and long-term interest rates. It has been one of the most widely tested propositions in the economics and finance literature.

Essentially, the EH states that future expected interest rates are implied by the current term structure. The pure expectations hypothesis states that one-period returns on a long-term bond and a short-term bond should be equal. This means that future expected interest rates are equal to forward rates implied by the current term structure. Alternatively, the return on holding a long-term bond to maturity should be equal to the expected return on investment in a series of short-term bonds over the life of the long-term bond.[12] The pure form of EH requires risk-neutrality on part of investors, who must be indifferent between investing in a long-term bond and rolling over a sequence of short-term bonds. That is, they require no premium when investing in a long-term bond.

Another theory that adds to the explanation of the term structure is the liquidity preference hypothesis proposed by Hicks (1946). It assumes that investors are risk averse and thus prefer the less risky short-term securities; they could be induced to hold longer-term,

[12] This statement is true in the case of certainty. For uncertain interest rates, deviations between long- and short-bond returns should follow a zero-mean white noise process.

more risky, securities if offered extra yield. The difference in yields has been labeled a term premium. The amount that constitutes the term premium is expected to increase with maturity because longer-term debt securities have higher price risk (an increasing default risk can also be of concern if the security is not default risk-free).

The liquidity preference theory does not contradict the EH, only its pure form. It is possible to test empirically whether the EH holds by allowing for the term premium in the relation between short and long interest rates. However, the term premium must be constant for a given maturity for the EH to hold.

The market segmentation hypothesis proposed by Gilbertson (1957) offers a different explanation of term premia. Market participants are assumed both to be risk averse and to have strong maturity preferences. They would not invest in bonds of adjacent maturities even if offered a premium. Thus, even bonds of close maturities are not substitutes of each other. Recognizing the excessive restrictiveness of the segmentation hypothesis, Modigliani and Sutch (1966) proposed the preferred habitat theory (PH). According to the PH, investors who for some reason prefer certain maturities to others may still be induced to invest in other maturities if offered a sufficiently large premium.

Numerous empirical tests of the expectations hypothesis in its various forms have been undertaken. While the literature review that follows will mention some of these studies, by no means its purpose is to provide an extensive coverage of the literature; this task would require a separate paper to be accomplished. Overall, empirical studies have rejected the expectations hypothesis more often than they have failed to do so. Also, the EH has been rejected more often at the short end of the term structure than for longer maturities. Different explanations (discussed later) have been proposed for the failure of the EH. To our

knowledge, only one paper (Brown et al. (2004)) has attempted to combine the EH and PH in empirical tests.

The purpose of this study is to test whether the EH holds at the short end of the term structure when preferred habitat is accounted for. The same data will be used as in the previous study where we found that the four major world currencies – USD, Euro, JPY, and SF – have economically and statistically significant quarter-end and year-end effects consistent with the preferred habitat for liquidity in the one-week and one-month LIBOR. In the light of our previous findings, preferred habitat in the shortest maturities may contribute to the failure of the EH at the short end of the term structure.

2.2 Literature Review

Modigliani and Sutch (1966) argue that preference for consumption at certain points in time creates “habitats”. Investors can be induced to leave their natural habitats if offered higher expected returns. Cox, Ingersoll, and Ross (CIR, 1981) argue that it is not preference for consumption at different times but the degree of risk aversion that creates liquidity habitats. They also show that under stochastic interest rates different term premia cannot simultaneously equal zero because of Jensen’s inequality. This implies that different versions of the pure expectations theory (which states that all term premia must simultaneously equal zero) are inconsistent with one another. Inconsistency in this context implies existence of arbitrage opportunities. However, Campbell (1986) demonstrates that different versions of the expectations theory, as opposed to the pure expectations theory, may be compatible with each other or arbitrage pricing equilibrium. The premia may not equal zero, but may still be

constant over time. He also argues that the differences among term premia are second-order effects of bond yield variability and are negligible in empirical tests of the theory.

McCuloch (1993) and Fisher and Gilles (1998) present some counterexamples to the CIR proof. They show that traditional forms of the EH can be consistent with the absence of arbitrage. They acknowledge, however, that their results are economically implausible.

Longstaff (2000a) notes that the analysis of CIR (1981) was developed under the assumption of complete markets; thus, bonds are redundant securities. Traditional forms of the EH can hold without arbitrage in a more realistic case where bond prices exhibit security-specific variation. This makes the validity of the EH an empirical issue; the EH cannot be defied on theoretical grounds.

Cook and Hahn (1990) provide a survey of the literature on the determinants of the yield curve. The studies surveyed in their paper find that over long periods of time the yield curve from three to twelve months has had negligible power to predict interest rates three to six months into the future. The research in this area suggests that the poor forecasting power of the yield curve from three to twelve months is due to substantial variation in the term premium at the three- and six-month horizons.

Hardouvelis (1988), Simon (1990) and Roberts, Runkle and Whiteman (1996) test the expectations hypothesis using the effective federal funds rate as the short-term rate and the three-month T-bill rate as the long-term rate. They reject expectations with two exceptions. First, all of them find that the EH holds during the period over which the Fed targeted nonborrowed reserves, October 1979 through October 1982. The second exception occurs when Roberts, Runkle and Whiteman (1996) test the EH using only settlement Wednesdays. Settlement Wednesday is the last day of the reserve requirement period when banks are

required to meet reserve requirements imposed by the Fed. Large spikes in the federal funds rate have often occurred on settlement Wednesdays as banks attempt to meet their reserve requirements. Both results seem counterintuitive in the light of EH being a proposition about predictability of the short-term rate. Indeed, the federal funds rate should be less predictable both when the Fed targets monetary aggregates as opposed to the federal funds rate and on settlement Wednesdays.

Thornton (2004a) shows that the counterintuitive results of Roberts, Runkle and Whiteman (1996) and other authors are primarily due to the fact that the short-term rate appears in both sides of the regression equation utilized in the tests. He points out that the results favoring the EH in the case of settlement Wednesday tests are generated because large changes in the federal funds rate tend to occur on these days. The results seemingly supporting the EH under the monetary aggregate targeting are due to the fact that the federal funds rate rose relative to the T-bill rate when interest rates were rising and fell relative to the T-bill rate when interest rates were falling. While the exact reason for this is not clear, it created a counter-cyclical pattern in the behavior of the spread between the short and long rates. That is, the spread between the three-month T-bill rate and the federal funds rate tended to widen when interest rates rose and narrow when interest rates declined. It caused, in turn, the failure to reject the EH, and not the T-bill rate adjusting in anticipation of funds rate movements.

Bekaert, Hodrick, and Marshall (1997) argue that regression-based tests of the expectations hypothesis tend to reject it because of biases that arise in small samples in such estimations. The biases are primarily due to high persistence in short interest rates. The

authors propose to use bias-adjusted VAR-GARCH models to test the expectations hypothesis for short-term rates.

Hurn, Moody, and Muscatelli (1995) conduct tests of the EH using monthly LIBOR data for the Pound Sterling over the period January 1975 through December 1991. Using cointegration analysis and vector autoregressions, they find relatively strong support for the expectations hypothesis. Cuthbertson (1996), who utilizes weekly data for GBP LIBOR (maturities of one week, one, three, six, and twelve months), finds some support for the EH except for the six-twelve month pair of maturities.

Longstaff (2000b) tests the expectations hypothesis at the extreme short end of the term structure using U.S. repurchase agreement (repo) rates over the period May 21, 1991 through October 15, 1999. Using these data, he tests the following parameterization of the expectations hypothesis:

$$\frac{1}{k}\sum_{t}^{t+k-1} r^{m} - r_t^{n} = a + \beta r_t^{n} + \varepsilon_t \ ,$$

where r^m is the short (m-period) rate, r^n is the long (n-period) rate, and $k = n/m$ is an integer.

In Longstaff (2000b), an overnight repo rate was selected the short rate, while the long rates included one-week, two-week, three-week, one-month, two-month, and three-month repo rates. The intercept of this model is a constant term premium that need not be equal across different maturities *n*. In the case when the pure version of the expectations hypothesis holds, the term premium must not be significantly different from zero. Under the expectations hypothesis, b_n also should be indistinguishable from zero. Longstaff (2000b) finds that the pure expectations hold for the repo rates over the period of his analysis. That is,

estimates of a_n and b_n are numerically small and not statistically different from zero. He also runs the estimations using the bias-adjusted VAR-GARCH model proposed by Bekaert, Hodrick, and Marshall (1997) and arrives at the same conclusion.

Brown, Cyree, Griffiths, and Winters (2004) decide to reexamine Longstaff's (2000b) results because they find his result puzzling for the U.S. repo market which is known to have a year-end preferred habitat (Griffiths and Winters (1997)). Brown et al. first replicate the results of Longstaff for the same data period using both OLS and bias-adjusted VAR-GARCH model. Their conclusion coincides with the Longstaff's: the pure expectations hypothesis holds for every term repo spread over the overnight repo rate over the period 5/91 to 10/99. To test whether Longstaff's results may be generalized to other time periods, Brown et al. conduct the same set of procedures on out-of-sample data (relative to the Longstaff sample). The data period is from February 1984 to May 1991, which predates the Longstaff's. Repo rate levels, volatility, and term premiums were much higher during this period compared to the later sample period used by Longstaff (2000b). The expectations hypothesis is soundly rejected for every term in the out-of-sample tests. A closer look at the repo data for the period from 1960 through 2003 suggests that the out-of-sample data may be more representative of normal market conditions. Analysis of the data over the period that follows the Longstaff sample period yields results similar to those of Longstaff, which may imply a new interest rate environment. Tests of the EH using daily LIBOR data for the four of the major world currencies (USD, Euro, JPY, and AUD) over the period from January 2001 through December 2002 yield mixed results: the JPY data reject the expectations hypothesis. The overall conclusion of Brown et al. (2004) is that Longstaff (2000b) results are sample-specific.

Downing and Oliner (2004) test the expectations hypothesis in the U.S. commercial paper (CP) market. Their data are based on actual transaction prices and cover the period from January 1998 through August 2003. Prior to 1998, the Fed collected quotes from commercial paper dealers, which the authors suggest might be inferior to transaction-based data. Downing and Oliner note that the CP market is characterized by large yield increases at the end of the year. They attribute these increases to either window dressing by financial institutions, as in Musto (1997), or the desire of CP issuers to lock in longer-term interest rates and avoid highly volatile overnight rates immediately prior to the turn of the year. Downing and Oliner do not consider the year-end preferred habitat for liquidity as a possible explanation of the year-end effect in the CP market. Downing and Oliner report that the year-end increase in term premia (spreads between term CP and one-day CP) causes the expectations hypothesis to be rejected. However, when the year-end effect is controlled for, the results are much more supportive of the expectations hypothesis. The dealer-quoted data collected prior to 1998, however, reject the expectations hypothesis even after controlling for the year-end effect (although the significance of term premia coefficients declines after the controls are added). The proposed explanation is that short-term interest rates have become more predictable since 1994 (as reported by Lange et al (2003)), although the authors do not rule out the lower quality of dealer-quote data contributing to this result.

According to the expectations hypothesis, a rise in the long rate relative to the short rate is due to the expectation of higher short rates in the future. If such predictions are on average correct, future short rates would tend to rise, creating a positive correlation of the changes in the short rate with the earlier spread. Campbell and Shiller (1991), among others, find for the U.S. that the spread predicts the wrong direction in the subsequent long rate changes: an

increase in the long-short spread is followed by a decline in the long rate next period. The authors find this behavior puzzling: while the movement of future cumulative short rates obeys the direction predicted by the expectations hypothesis, the short-run movement of long rates does not. Hardouvelis (1994) conducts the analysis for USD, GBP, JPY, CAD, DM, FF, and ITL using a ten-year rate as the long rate and a three-month rate as the short rate. He finds that in France and Italy the long rate moves in the correct direction. In Canada, Germany, Japan, and the U.K., the long rate moves in the opposite direction from the one predicted by the expectations hypothesis, but this movement is due to simple white noise and is reversed by the use of instrumental variables. In the U.S., however, white noise cannot explain the puzzle. The only hypothesis that Hardouvelis considers feasible is the overreaction hypothesis stating that short-term interest rates change too much on monetary policy announcements and that this error is subsequently corrected. Tzavalis and Wickens (1997), among others, attribute the Campbell-Shiller paradox to time-varying term premia.

Thornton (2004b) provides an econometric resolution to the seemingly paradoxical results of Campbell and Shiller (1991). He shows that the tests of the EH that are the most popular in empirical work by construction generate results consistent with those of Campbell and Shiller when the EH does not hold, whatever the reason. Monte Carlo experiments conducted by Thornton support this explanation.

Patel and Shoesmith (2004) analyze integration among interest rates in Germany, Japan, the U.K., and the U.S. and find that the one-month rate generally has little or no influence in setting the long-term trend. The authors interpret this result as the consequence of the one-month rate being much more influenced by monetary policy changes than longer term rates. While this may be true, in the context of this study, quarter-end and year-end preferred

habitat for liquidity in the short-term maturity may also contribute to the lack of influence of the shortest rates on long-term trends.

In this study, due to rather short data availability periods for very short-term LIBOR (maturities below one month), the one-month LIBOR has been selected the short-term rate for this study. The long-term interest rates are three-month, six-month, and twelve-month LIBOR. Of course, even the longest maturity here would still represent a short-term loan when it comes to the traditional classification used by capital market participants.

2.3 Methods and Empirical Results

2.3.1 Preliminary Empirical Tests Using OLS

Until recently, OLS has been a traditional framework for testing the EH. It is well known, however, that OLS often does a poor job with time series data, particularly with interest rates, due to high persistence in the data. Lately, the literature has been leaning toward the VAR-based tests of the EH proposed by Bekaert et al. (1997) and Bekaert and Hodrick (2001). Acknowledging the shortcomings of OLS, we use it as a starting point for our analysis. The EH posits that the long-term rate is determined by the market's expectations of the short-term rate over the life of the long-term rate. To test this assertion, parameterizations of algebraic manipulations of the following equation have been used:

$$r_t^n = (1/k)\sum_{i=0}^{k-1} E_t r_{t+mi}^m + \pi^{n,m} \qquad (2),$$

where r_t^n is the n-period (long-term) rate, r_t^m is the m-period (short-term) rate, $\pi^{n,m}$ is the constant risk premium that may vary with maturity of the rates, and k = n/m is an integer.

One frequently used parameterization of (2) is

$$\bar{r}^m - r_t^m = \alpha + \beta(r_t^n - r_t^m) + \varepsilon_t \qquad (3),$$

where $\bar{r}^m$ is the average short rate over the time span of the long rate. If the EH holds, beta will be indistinguishable from one, that is, the spread between the long and short rates will not have predictive power for the future short-term rate behavior. If the pure form of EH holds, the intercept in (3) must be indistinguishable from zero. The pure EH implies flat term structure.

A different OLS test of the EH derived from (2) was used by Longstaff (2000b):

$$\bar{r}^{m} - r_t^n = \alpha + \beta r_t^n + \varepsilon_t \qquad (4).$$

Here both the intercept and beta must not be different from zero for the pure version of the EH to hold; beta must be indistinguishable from zero for the EH to hold.

We start with running equations of the form (3) and (4) for different currencies. Since our goal is to test for the expectations hypothesis in the markets with identified preferred habitat (PH) for liquidity, we continue by adding two PH dummy variables, QPH and YPH, to (3) and (4). Recall that the spread between one-month LIBOR and longer-term LIBOR for USD, Euro, SF, JPY, DM, and ITL increases significantly on the second to last day before the maturity of the loan crosses the end of the year and is sustained throughout the third to last day of the year. This year-end increase is also statistically significant in GBP, but it is small in magnitude. The quarter-end preferred habitat effects in a one-month maturity are much smaller; they are significant only for USD and JPY. To account for this behavior in the one-month LIBOR, we create two dummy variables: QPH, which covers a period of the increased spreads preceding the quarter-end, and YPH, covering a similar period prior to the year-end. For example, in 1989 QPH for the second quarter is set to one for the period May 30 through June 28, while YPH in the same year is equal to one on each trading day from November 29 through December 27. We separate year-ends from other quarter-ends because the turn-of-the-year effect in the one-month maturity was found to be much more significant, both statistically and economically, than the turn-of-the-quarter effect. After adding the PH dummies to the regression model, (3) looks as follows:

$$\bar{r}^{m} - r_t^m = \alpha + \beta(r_t^n - r_t^m) + C_1 QPH + C_2 YPH + \varepsilon_t \qquad (3')$$

Given our earlier findings, C_2 is expected to be negative for USD, Euro, SF, JPY, DM, and ITL, while C_1 is also expected to be negative for USD and JPY. Also, C_1 is expected to be smaller in absolute value than C_2. The negative sign is expected because the LHS of (3') should decrease during the period preceding the end of the quarter or year due to the increase in the one-month rate, r_m.

Our task is to determine whether the preferred habitat dummies belong to the model and whether they have a significant influence on the relation between the short and the long rate. For example, the EH may not hold without the PH variables in the model, but may hold when they are included. Table 6 contains output for equations 3 and 3' with the three-month LIBOR as the long rate. In order for the EH to hold, the coefficient of the spread between the long and the short interest rates must not be different from unity. Thus, p-values of these coefficients (betas) reported in Table 6, are for the null hypothesis $\beta = 1$.

Table 6. The traditional model of the Expectations Hypothesis with and without preferred habitat variables.

Without PH variables			With PH variables			
Currency	**Alpha**	**Beta**	**Alpha**	**Beta**	**QPH**	**YPH**
USD	**-0.093**	0.874	**-0.069**	**0.774**	-0.014	**-0.159**
	0.000	0.179	**0.000**	**0.017**	0.399	**0.000**
GBP	**-0.086**	**0.799**	**-0.086**	**0.800**	-0.004	0.013
	0.000	**0.021**	**0.000**	**0.021**	0.897	0.686
AUD	**-0.084**	0.955	**-0.077**	0.954	-0.025	-0.015
	0.000	0.619	**0.000**	0.607	0.243	0.529
DM	**-0.036**	**0.593**	**-0.024**	**0.558**	-0.007	**-0.108**
	0.000	**0.000**	**0.030**	**0.000**	0.778	**0.005**
Euro	**-0.052**	**0.733**	**-0.039**	**0.706**	-0.011	**-0.118**
	0.000	**0.013**	**0.006**	**0.008**	0.631	**0.001**
SF	**-0.066**	**0.687**	**-0.050**	**0.642**	-0.022	***-0.100***
	0.000	**0.002**	**0.003**	**0.000**	0.413	***0.090***
JPY	-0.025	0.927	**-0.019**	**0.902**	-0.021	-0.016
	0.182	0.127	**0.008**	**0.044**	0.110	0.435
CAD	**-0.131**	0.932	**-0.134**	0.929	0.017	-0.008
	0.000	0.635	**0.000**	0.622	0.689	0.868
FF	**-0.070**	1.012	-0.048	1.016	-0.077	-0.038
	0.015	0.932	0.139	0.910	0.203	0.710
SP	0.001	1.155	0.013	1.155	0.001	-0.049
	0.731	0.168	0.720	0.167	0.982	0.396
ITL	-0.008	0.955	0.035	0.942	-0.096	**-0.247**
	0.850	0.709	0.491	0.638	0.213	**0.038**

The short-term rate is a one month LIBOR, the long-term rate is a three-month LIBOR.
P-values based on Hansen-Hodrick standard errors are below the coefficients.
The p-values for betas are for the Wald test with the null hypothesis $\beta = 1$.
Coefficients significant at the 5% level or better are bold; those significant at the 10% level but not at the 5% level are bold and italicized.

According to the output of the model without the PH variables, the EH does not hold for four out of eleven currencies: GBP, Euro, SF, and DM. It holds for USD, which contradicts numerous results from the empirical tests of the expectations hypothesis. When included into the model, the year-end preferred habitat dummy is significant for USD, Euro, SF, DM, and ITL, while the quarter-end PH dummy is never significant, although its sign is in the expected direction. The insignificance of the turn-of-the-quarter dummy variable is understandable: the quarter-end PH effect is small and often statistically insignificant for the

one-month maturity. Also, when the PH dummies are included into the model, spread coefficients of the currencies with pronounced year-end effects deviate further from the theoretical value of one, while the intercepts (which reflect the term premia) become smaller in absolute value. This further deviation of the spread coefficients from unity leads to the rejection of the EH for two more currencies, USD and JPY (even though the PH variables by themselves are not significant for JPY, which is surprising).

According to the output, the EH holds for SP and ITL in its pure form, i.e., with the intercept and the coefficient of the spread not being different from zero and one, respectively. For AUD, CAD, and FF, the intercepts are different from zero thus rejecting the pure form of the EH, but the coefficients of the spread are indistinguishable from one. For these currencies, the coefficients of the PH dummies are insignificant, and the presence of these variables, as one would expect, does not change the nature of the relation between one-month and three-month rates, except for FF (the pure form of the EH holds after adding the PH dummies), which is unexpected.

Longstaff (2000b) estimated the model of the form (4). The long rate is subtracted on the LHS from the short rate averaged over the term spanned by the long rate. Under this setup, both alpha and beta must be equal to zero for the pure EH to hold. The estimation results are in Table 7. Again, the model was estimated with and without PH dummy variables. When estimation is done with the PH controls, equation (4) becomes

$$\overline{r}^{m} - r_t^n = \alpha + \beta r_t^n + C_1 QPH + C_2 YPH + \varepsilon_t \quad (4').$$

Table 7. The Longstaff (2000b) Test of the Expectations Hypothesis

Without PH variables			With PH variables			
Currency	**Alpha**	**Beta**	**Alpha**	**Beta**	**QPH**	**YPH**
USD	**-0.063**	***-0.007***	**-0.055**	***-0.007***	0.000	**-0.103**
	0.001	***0.087***	**0.006**	***0.086***	0.994	**0.003**
GBP	**-0.103**	0.001	**-0.104**	0.001	-0.002	0.021
	0.000	0.895	**0.000**	0.891	0.938	0.560
AUD	0.022	**-0.015**	0.030	**-0.015**	-0.025	-0.019
	0.378	**0.000**	0.224	**0.000**	0.217	0.317
DM	0.002	**-0.010**	0.000	**-0.012**	0.010	-0.042
	0.950	**0.018**	0.946	**0.020**	0.614	0.281
Euro	-0.027	-0.005	-0.019	-0.005	-0.003	**-0.085**
	0.249	0.305	0.403	0.298	0.890	**0.013**
SF	**-0.090**	0.002	**-0.084**	0.002	-0.004	-0.049
	0.000	0.768	**0.000**	0.779	0.883	0.391
JPY	**-0.029**	-0.002	**-0.026**	0.002	-0.012	-0.002
	0.000	0.499	**0.000**	0.498	0.334	0.921
CAD	***0.056***	**-0.036**	***0.053***	**-0.036**	0.013	-0.009
	0.056	**0.000**	***0.058***	**0.000**	0.761	0.830
FF	-0.013	-0.008	0.010	-0.008	-0.077	-0.036
	0.833	0.432	0.878	0.431	0.198	0.732
ESP	-0.088	0.009	-0.083	0.009	-0.010	-0.056
	0.373	0.502	0.409	0.498	0.868	0.376
ITL	-0.142	0.014	-0.110	0.015	-0.089	**-0.248**
	0.482	0.555	0.577	0.524	0.249	**0.050**

The short-term rate is a one month LIBOR, the long-term rate is a three-month LIBOR.
P-values based on Hansen-Hodrick standard errors are below the coefficients.
Coefficients significant at the 5% level or better are bold; those significant at the 10% level but not at the 5% level are bold and italicized

The results reported in Table 7 are quite different from those in Table 6. For example, the EH now holds for GBP and Euro, but does not hold for AUD. This suggests that the estimation results are dependent on the parameterization, and that the OLS likely produces biased estimations. One similarity with the output from The year-end PH dummy is statistically significant for USD, Euro, and ITL.

The same analyses were conducted with the six-month and twelve-month LIBOR as the long rate. Again, conclusions differ depending on the model used. It suggests that OLS may

yield biased estimates and that more appropriate techniques should be used to draw better conclusions regarding the EH and the influence of PH on it.

The regressions whose outputs are reported in Tables 6 and 7 suffer from very low R-squareds and Durbin-Watson statistics close to zero. It suggests high autocorrelations in residuals which, if uncorrected, may lead to biased estimates. In the OLS framework, one of the following methods may be applied to correct this problem: Cochrane-Orcutt, Hildreth-Lu, and maximum likelihood. All three methods were used, and the outputs were then compared. Under any of these methods, R-squareds become high (close to unity), and Durbin-Watson statistics are around two. However, the data are not fitted well in that intercepts and sometimes betas become very large in absolute values. Although the year-end PH dummy is significant for each of the "usual suspects" – USD, Euro, SF, and JPY – the results appear to lack credibility, and thus are not reported.

2.3.2 Tests of Cointegration

Cointegration between variables implies that they share common stochastic trends. In order for nonstationary variables to have a stable long-term relationship, they must be related through at least one linear set of coefficients (a cointegrating vector) that makes the combination of the variables stationary. Interest rates are usually assumed to follow an I(1) process. That is, the rate series are assumed to be nonstationary with a unit root; taking the first difference is needed to obtain stationary series.[13] The augmented Dickey-Fuller test results suggest that this is mostly true for our data, as interest rate series for ten out of eleven

[13] BBA reports annualized interest rates. For all subsequent tests, we use continuously compounded equivalents of the annualized rates given by $r_c = (12/m)*ln(1+r*(m/12))$, where r is the annualized rate, and m is maturity in months.

currencies are nonstationary with a unit root, while the null of non-stationarity can be decisively rejected for the first differences of these series. The only exception is CAD, for which the null hypothesis of the unit root can be rejected at the 5% level for all four maturities. (Results of these tests are not reported for brevity.)

Tests of cointegration between the long and the short interest rates may shed some light on the nature of the relationship between them. Thornton (2004c) notes that the lack of cointegration may suggest that EH is likely to fail. First, if the interest rates are truly integrated of order one (I(1)), rejecting the hypothesis of cointegration implies that there is no stable relationship between the levels of interest rates, and the EH cannot hold. However, the power to reject the null hypothesis of nonstationarity is low when the root is close to 1. That is, it could be that interest rates are actually I(0). If it is true, however, it should not be difficult to find evidence of cointegration, i.e., reject the null hypothesis that there is no stationary relationship between the short-term and long-term rates. Thus, failure to find at least one cointegrating relationship among stationary variables is relatively strong evidence against the EH.

Because there are only two variables in our tests (the long rate and the short rate), the maximum possible number of cointegrating vectors is one.[14] If the rates are nonstationary but cointegrated, one can test the EH by testing the null that the cointegrating vector equals (1, -1) after adjusting for the constant risk premium and/or a deterministic trend). As Hall, Anderson, and Granger (1992) point out, the cointegrating vector of (1, -1) implies that the spread between long- and short-term interest rates ($1^*R_L - 1^*R_S$) is the stationary linear combination between them which results from cointegration. Rejecting the hypothesis that

[14] Chapter 6 in Enders (1995) provides an excellent treatment of the topic of cointegration

the cointegrating vector is (1, -1) suggests that the long-run equilibrium relationship is in the direction inconsistent with the EH.

As Thornton (2004c) notes, if the EH does not hold in the long run, it is unlikely to hold at frequencies that are of interest to policymakers or financial analysts. On the other hand, finding that the EH holds in the long run does not automatically imply that the EH holds at higher frequencies. For that to happen, the long-term rate must respond rather quickly to changes in the policy rate. That is, failing to reject that the cointegrating vector is (1, -1) does not necessarily mean that the EH holds at frequencies that are of interest to policymakers and financial analysts.

Having found that interest rates in our sample are integrated of order one (I(1)), we proceed by testing for a unit root in the spread between the long and short rates. As noted above, for the EH to hold in the long run, it is necessary that the spread between the rates be stationary. Thus, the rejection of the unit root for the spread between long and short rates is a necessary condition for the EH to hold in the long run. Table 8 contains the augmented Dickey-Fuller (ADF) test statistics for the spreads between long and short interest rates for each currency. The number of lags in the ADF test was chosen using the Akaike information criterion. The null hypothesis of a unit root was rejected at the 5% confidence level for all but two spreads. For the spreads between the twelve-month and one-month LIBOR in DM and AUD, the null of a unit root could only be rejected at the 10% confidence level. The strength of cointegrating relations (expressed by the magnitude of the ADF statistics) monotonically declines with maturity of the long rate. This result is similar to Thornton (2004c), who finds weaker and eventually insignificant relations between the yields of three-month and longer term Japanese securities as maturity of the long rate increases. Overall, the long and short

rates are cointegrated within each currency in our sample, which implies that they share common stochastic trends, while the spreads are stationary. It is consistent with the expectations hypothesis holding in the long run; however, it does not necessarily mean that the EH holds at frequencies that may be of interest to policymakers and practitioners.

Table 8. Augmented Dickey-Fuller tests of for spreads between one-month LIBOR and longer term rates.

Currency/Long rate	3 months	6 months	12 months
USD	**-9.04**	**-6.44**	**-4.80**
GBP	**-6.91**	**-4.92**	**-3.34**
AUD	**-6.30**	**-4.03**	***-2.70***
Euro	**-8.69**	**-5.83**	**-4.14**
SF	**-7.73**	**-5.65**	**-3.86**
JPY	**-7.99**	**-5.63**	**-4.25**
CAD	**-7.56**	**-5.62**	**-4.45**
DM	**-5.80**	**-3.66**	***-2.59***
FF	**-7.40**	**-5.59**	**-4.47**
SP	**-8.72**	**-6.91**	**-5.81**
ITL	**-7.24**	**-5.50**	**-3.20**

Note: Dickey-Fuller statistics significant at the 5% level or better are bold; those significant at the 10% level are bold and italicized.

Rodrigues and Franses (2003), among others, raise a concern that unit root tests of low-frequency data, such as daily or weekly, may have relatively low power, that is, a fairly high chance of accepting the null of the unit root when it is false. It is not a major concern with our data sample since the null of the unit root for spreads has been rejected uniformly across currencies at traditional confidence levels. As to the unit root tests for interest rate levels, the null of the unit root cannot be rejected for any of the currencies and maturities even at extremely conservative levels of confidence.

Another way to see if interest rates may be cointegrated is to run the Granger causality tests. Variable x is said to Granger cause variable z when lags of x enter into the equation for z. Granger causality does not imply that one variable causes the other, although it may. The

absence of Granger causality, however, implies no cointegrating relations between the variables. At least one of them must Granger cause the other if the two variables are cointegrated. We find that longer rates for all currencies Granger cause one-month rates, with no exception. As to the one-month rates, we only find two exceptions: the one-month JPY LIBOR does not Granger cause any of longer JPY rates, while the one-month SF LIBOR does not Granger cause the twelve-month SF LIBOR. Thus, Granger causality test results suggest that one-month LIBOR may be cointegrated with longer term rates within each currency.

Another weak test of the EH is that the spread between long- and short-term rates Granger causes changes in the short rate. This hypothesis could not be rejected at the 1% level for any of the currencies or maturities when tested.[15]

To further examine whether the relations between short and long rates are consistent with the EH, it is useful to estimate cointegrating vectors. A cointegrating vector is a certain set of non-zero coefficients that makes a combination of variables stationary. Estimated cointegrating vectors can shed some light on the nature of the long-term relation between the variables. As mentioned above, in order for the EH to hold, the relation between short- and long-term rates must be expressed by the cointegrating vector equal to (1, -1), that is, the spread between the rates must be stationary. We already found with the help of the ADF test that spreads between long and short interest rates were stationary. However, the Johansen test (based on the works of Johansen (1991, 1995)) is typically used to estimate the number of cointegrating vectors and the vectors themselves for each combination of rates. In the case of two variables, there may be at most one cointegrating vector. Since any linear combination

[15] The output of the Granger causality tests is not reported to save space.

that is a multiple of (1, -1) would also be a cointegrating vector (e.g., (2, -2) or (0.1, -0.1)), it is convenient to normalize the vector with respect to one of the variables. For this purpose, the coefficient of the short-term rate was set equal to one (nothing of substance would change if the coefficients were normalized with respect to the long rate). The output of the Johansen procedure can then be used to test (using a chi-square distribution) whether the coefficient of the long rate is significantly different from -1 in the cointegrating equation.

The Johansen tests were performed for 33 pairs of interest rates, three pairs in each of the eleven currencies (one-month LIBOR vs. three-, six-, and twelve-month LIBOR). The test output suggests that the short rate (one-month LIBOR) is cointegrated with longer rates (three-, six-, and twelve-month LIBOR for the same currency) for every currency with exactly one cointegrating vector. According to the likelihood ratio test statistic, the cointegrating relations are significant at the 1% level for most of the interest rate pairs, with the rest of them being significant at the 5% level. Existence of cointegration between short and long rates is expected; this result just re-confirms the conclusion of the ADF test which could not reject stationarity for the spreads. It implies that there is a long-term relationship between long and short interest rates. The number of lags for the Johansen test was chosen using the Akaike information criterion. The number of lags does not seem to affect the strength and direction of relations between short-term and long-term interest rates in any of the 33 pairs. Table 9 presents the output of the Johansen test for all eleven currencies.

Table 9. Cointegrating coefficients of the long-term rates.

Long rate	3m	6m	12m
Currency			
USD	-0.990 (0.002)	-0.982 (0.005)	-0.963 (0.011)
χ^2	16.158 [0.000]	12.073 [0.001]	11.422 [0.001]
GBP	-1.000 (0.002)	-1.009 (0.006)	-1.020 (0.013)
χ^2	0.000 [0.984]	2.427 [0.119]	2.362 [0.124]
AUD	-1.001 (0.002)	-1.002 (0.005)	-0.987 (0.014)
χ^2	0.495 [0.482]	0.120 [0.729]	0.892 [0.345]
Euro	-1.002 (0.003)	-1.007 (0.008)	-1.016 (0.015)
χ^2	0.271 [0.603]	0.751 [0.386]	1.088 [0.297]
SF	-0.999 (0.004)	-1.011 (0.010)	-1.032 (0.024)
χ^2	0.029 [0.865]	1.204 [0.272]	1.840 [0.175]
JPY	-1.003 (0.003)	-1.014 (0.006)	-1.026 (0.012)
χ^2	1.041 [0.308]	4.711 [0.030]	4.629 [0.031]
CAD	-0.990 (0.003)	-0.973 (0.009)	-0.936 (0.017)
χ^2	8.623 [0.003]	10.054 [0.002]	14.145 [0.000]
DM	-0.996 (0.004)	-0.999 (0.010)	-1.010 (0.021)
χ^2	0.755 [0.385]	0.018 [0.894]	0.240 [0.624]
FF	-1.004 (0.004)	-1.018 (0.011)	-1.033 (0.021)
χ^2	0.895 [0.344]	2.640 [0.104]	2.438 [0.118]
SP	-1.020 (0.004)	-1.036 (0.009)	-1.050 (0.016)
χ^2	23.950 [0.000]	15.148 [0.000]	9.608 [0.002]
ITL	-1.019 (0.007)	-1.030 (0.013)	-1.032 (0.021)
χ^2	8.390 [0.004]	5.115 [0.024]	2.356 [0.125]

Coefficients of the short-term (one-month) interest rate are set to equal one. Estimated coefficients of the long rate are in the table. Standard errors are in parentheses next to the coefficients. The null hypothesis is that the coefficient of the long rate is equal to -1. The chi-square statistics testing the null are below coefficients. They p-values for the chi-square test are to the right of the chi-squire statistics, in square brackets.

As is evident from the output reported in Table 9, all estimated cointegrating vectors are quite close to -1. However, the null hypothesis that a particular vector is equal to (1, -1) is sometimes rejected due to small standard errors.[16] Still, the cointegrating vectors are not different from (1, -1) for all three pairs of rates in six out of eleven currencies (GBP, AUD,

[16] The form of a cointegrating equation with an intercept was also considered. It leads to different conclusions sometimes: for example, the long rate coefficients of USD become indistinguishable from -1, while those of GBP and AUD become different from -1. We do not report these results because we believe that the cointegrating vectors should be estimated without the intercept: the intercepts implies a deterministic trends in the data. While interest rates increase or decrease during some periods, no trend persists throughout the entire data period for any of the currencies. However, given dependence of the results on form of the cointegrating vector, caution should be exercised when interpreting the output from Table 9.

Euro, SP, DM, and FF). The coefficients of the long-term rate are close to, although often not statistically indistinguishable from, -1 in the rest of the interest rate pairs.

Because this effectively is a joint test of the exact I(1) behavior and the EH, it should not be taken as a serious rejection of the EH. Although according to the Johansen tests not all cointegrating vectors are indistinguishable from (1, -1), recall that all the spreads were found to be stationary, which is consistent with the EH holding in the long run. On the other hand, finding that interest rates are cointegrated with the cointegrating vector indistinguishable from (1, -1) only implies that the long-run behavior of rates is consistent with the EH. Overall, we can conclude that the EH is likely to hold in the long run for pairs of interest rates in our sample. It does not necessarily mean that the EH holds at higher frequencies that are of interest for policymakers and analysts. To be useful for them, longer-term rates must respond quickly to changes in the market's expectation of the short-term rate. Definitely, a further investigation is warranted to determine whether it happens.

Next, we estimate the cointegrating equations controlling for the preferred habitat for liquidity before the end of the year. Variable YPH, defined as in equation (3'), is entered as the exogenous variable when cointegrating vectors are estimated. Overall, its influence is reflected in the coefficient of the long rate in the cointegrating vectors becoming smaller in absolute value. This is more noticeable in currencies with identified year-end preferred habitat for liquidity – USD, Euro, SF, JPY, and DM. The conclusion of the test, however, very rarely changes even for interest rates in these currencies. Cointegrating vectors for other

currencies are influenced less by the addition of the control variable for the year-end PH. The results are reported in Table 10.[17]

Table 10. Cointegrating coefficients of the long-term rates (after accounting for the year-end preferred habitat for liquidity).

Long rate	3m	6m	12m
Currency			
USD	-0.984 (0.002)	-0.973 (0.004)	-0.946 (0.010)
χ^2	64.860 [0.000]	41.929 [0.000]	29.990 [0.000]
GBP	-1.000 (0.002)	-1.008 (0.006)	-1.018 (0.013)
χ^2	0.041 [0.840]	1.782 [0.182]	1.805 [0.179]
AUD	-1.002 (0.002)	-1.003 (0.005)	-0.991 (0.014)
χ^2	0.664 [0.415]	0.294 [0.588]	0.467 [0.494]
Euro	-1.000 (0.003)	-1.005 (0.008)	-1.014 (0.015)
χ^2	0.001 [0.973]	0.409 [0.522]	0.839 [0.360]
SF	-0.991 (0.004)	-0.996 (0.010)	-0.995 (0.026)
χ^2	4.432 [0.035]	0.151 [0.697]	0.032 [0.859]
JPY	-1.003 (0.003)	-1.010 (0.006)	-1.020 (0.012)
χ^2	0.097 [0.755]	2.769 [0.096]	2.732 [0.098]
CAD	-0.990 (0.003)	-0.972 (0.009)	-0.935 (0.018)
χ^2	8.583 [0.003]	9.803 [0.002]	13.850 [0.000]
DM	-0.990 (0.004)	-0.986 (0.009)	-0.989 (0.019)
χ^2	7.239 [0.007]	2.284 [0.131]	0.325 [0.569]
FF	-1.004 (0.005)	-1.016 (0.012)	-1.025 (0.022)
χ^2	0.801 [0.371]	1.903 [0.168]	1.262 [0.261]
SP	-1.020 (0.004)	-1.035 (0.010)	-1.048 (0.017)
χ^2	22.477 [0.000]	13.850 [0.000]	8.447 [0.004]
ITL	-1.017 (0.007)	-1.028 (0.014)	-1.030 (0.021)
χ^2	6.281 [0.012]	4.180 [0.041]	1.949 [0.163]

Coefficients of the short-term (one-month) interest rate are set to equal one. Estimated coefficients of the long rate are in the table. Standard errors are in parentheses next to the coefficients. The null hypothesis is that the coefficient of the long rate is equal to -1. The chi-square statistics testing the null are below coefficients. They p-values for the chi-square test are to the right of the chi-squire statistics, in square brackets.

In sum, the analysis of cointegration suggests that interest rates are integrated of order one (that is, they are nonstationary (with the possible exception of CAD)), while first differences and spreads are stationary. One-month LIBOR is cointegrated with longer-term rates within each currency. Estimated normalized cointegrating vectors are sometimes

[17] Exogenous variable QPH, covering the quarter-end preferred habitat, had negligible effect on the estimation and always was insignificant. It is also true for all subsequent tests in the paper. Thus, we exclude it from estimations.

statistically different from the theoretical value of (1, -1) but they are always close to it. Combining this with the fact that spreads between one-month and longer term rates were all found to be stationary (or, more precisely, the null of non-stationarity has been rejected for each spread), we believe that there are solid grounds for the expectations hypothesis to hold in the long run. Whether it holds in the short run is the question that we attempt to answer next.

2.3.3 Tests of the Expectations Hypothesis Using VAR

Campbell and Shiller (1991) suggested that the EH could be tested using a general vector autoregression (VAR), and Bekaert and Hodrick (2001) made the procedure operational. Since a VAR encompasses a wider range of data generating processes than (2), VAR-based tests of the EH will be more powerful. The restrictions on the VAR implied by the EH can be tested using the procedure of Bekaert and Hodrick. They propose to utilize a constrained Generalized Method of Moments (GMM) estimation. Developed by Hansen (1982), GMM uses orthogonality conditions defined by the theory to develop tests. The orthogonality conditions are based on the assumption of rational expectations, under which a realization of a random variable is equal to its conditional expectation plus an error term which is orthogonal to the information that was used to form the expectations.

The VAR takes the form

$$(I - \Theta(L))y_t = \eta_t \qquad (5)$$

for y = (r_m, r_n)'; here L is the lag operator. GMM estimation imposes orthogonality conditions of the form $g(z_t, \theta) \equiv \eta_t \otimes x_{t-1}$, where x_{t-1} is a vector formed from stacking lagged

values of y, z_t is defined as $(y_t, x_{t-1})'$, and θ is a vector of parameters in Θ (L). Using the sample moment condition

$$g_T(\theta) \equiv \frac{1}{T}\sum_{t=1}^{T} g(z_t, \theta) \qquad (6),$$

GMM estimation proceeds by choosing θ to minimize the following objective function:

$$J_T(\theta) \equiv g_T(\theta)' W g_T(\theta) \qquad (7).$$

As Hansen (1982) shows, the optimal weighting matrix, W, is a consistent estimator of the inverse of

$$\Omega \equiv \sum_{k=-\infty}^{k=\infty} E[g(z_t, \theta) g(z_{t-k}, \theta)'] \qquad (8).$$

GMM is used to estimate restricted VARs by forming a Lagrangian from the usual GMM quadratic objective and a vector of parameter constraints. The Lagrangian is defined

$$L(\theta, \gamma) = -0.5 g_T(\theta)' \Omega_T^{-1} g_T(\theta - a_T(\theta)' \gamma \qquad (9),$$

where γ is a vector of Lagrange multipliers and the constraints on θ are represented by the vector valued function, $\alpha_T(\theta) = 0$. The matrix Ω_T is again a consistent estimate of the matrix Ω defined above. Denoting the Jacobian of $g_T(\theta)$ and $\alpha_T(\theta)$ by G_T and A_T, respectively, the first-order conditions for maximizing $\bar{\theta}$ and $\bar{\gamma}$ are

$$\begin{bmatrix} -G_T^{'} \Omega_T^{-1} \sqrt{T} g_T(\bar{\theta}) - A_T^{'} \sqrt{T} \bar{\gamma} \\ -\sqrt{T} \alpha_T(\bar{\theta}) \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix} \qquad (10).$$

The asymptotic distribution of the constrained estimator is derived from these first-order conditions. This is done by expanding $g_T(\theta)$ and $\alpha_T(\theta)$ around the true parameter value, θ_0, and substituting these into the first-order conditions above. This leads to a system of the form

$$\begin{bmatrix} -G_T^{'}\Omega_T^{-1}\sqrt{T}g_T(\theta_0) \\ 0 \end{bmatrix} - \begin{bmatrix} B_T & A_T^{'} \\ A_T & 0 \end{bmatrix} \begin{bmatrix} \sqrt{T}(\overline{\theta}-\theta_0) \\ \sqrt{T}\overline{\gamma} \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix} \qquad (11),$$

for $B_T \equiv G_T^{'}\Omega_T^{-1}G_T$. Use of the partitioned inverse formula allows one to argue that the constrained estimator, $\overline{\theta}$, is distributed as $\sqrt{T}(\overline{\theta}-\theta_0) \to N(0,\Sigma_T)$ for

$$\Sigma_T \equiv B_T^{-1} - B_T^{-1}A_T^{'}(A_T B_T^{-1} A_T^{-1})^{-1} A_T B_T^{-1} \qquad (12)$$

and the Lagrange multipliers are distributed asymptotically as

$$\sqrt{T}\overline{\gamma} \to N\left(0,(A_T B_T^{-1} A_T^{'})^{-1}\right) \qquad (13)$$

The standard errors implicit in (13) are autocorrelation and heteroskedasticity consistent. If the constraints have a noticeable impact on parameter estimation (which means that the EH does not hold), the estimated Lagrange multipliers should significantly differ from zero. On the other hand, if the EH does hold in the data, the constraints implied by the EH will not influence parameter estimates significantly, and the Lagrange multiplier will be indistinguishable from zero. The asymptotic distributions given above can be used to show that a test that the multipliers are jointly zero can be based on the statistic

$$\sqrt{T}\,\overline{\gamma}'\left(A_T B_T^{-1} A_T^{'}\right)\overline{\gamma} \qquad (14).$$

This Lagrange multiplier statistic is asymptotically distributed as χ^2 with l degrees of freedom, where l is the number of restrictions. Bekaert and Hodrick (2001) examine Wald, Lagrange multiplier, and distance metric tests and conclude that the Lagrange multiplier test does best when it comes to testing the null of the EH. It tends to slightly underreject the null, while the other two tests are prone to overrejection.

Maximizing the Lagrangian above may be computationally troublesome, so Taylor series approximation to $g_T(\theta)$ and $\alpha_T(\theta)$ can be used again to derive a constrained estimate with

similar asymptotic properties. Instead of expanding around the true value, θ_0, the current estimate of the true value, θ_i, is used to form a better approximation, θ_{i+1}. The approximations are $g_T(\theta_{i+1}) = g_T(\theta_i) + g_T(\theta_{i+1} - \theta_i)$ and $A_T(\theta_{i+1}) = A_T(\theta_i) + A_T(\theta_{i+1} - \theta_i)$; we can substitute them into the first-order conditions for maximization to derive the following iterative method:

$$\begin{bmatrix} -G_T^{'}\Omega_T^{-1}\sqrt{T}g_T(\theta_i) \\ -\sqrt{T}a_T(\theta_i) \end{bmatrix} - \begin{bmatrix} B_T & A_T^{'} \\ A_T & 0 \end{bmatrix} \begin{bmatrix} \sqrt{T}(\theta_{i+1} - \theta_i) \\ \sqrt{T}\gamma_{i+1} \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix}. \qquad (15)$$

The unconstrained VAR parameter estimates are used as initial conditions. The procedure is then iterated until the constraints are satisfied. Because the moment conditions for VAR estimation should be uncorrelated over time, Ω_T is estimated by

$$\Omega_T = \frac{1}{T}\sum_{t=1}^{T} g(z_t, \theta_U) g(z_t, \theta_U)' \qquad (16),$$

evaluating the moment conditions at the unconstrained VAR parameter estimates.

The constraints that the EH imposes on a VAR can be seen by writing the VAR in the first-order form:

$$\begin{pmatrix} r_t^m \\ r_t^n \\ r_{t-1}^m \\ r_{t-1}^n \\ \ldots \\ r_{t-k}^m \\ r_{t-k}^n \end{pmatrix} = \begin{pmatrix} \theta_1 & \theta_2 & \ldots & \theta_k \\ I & 0 & \ldots & 0 \\ 0 & I & \ldots & \ldots \\ \ldots & \ldots & \ldots & \ldots \\ 0 & \ldots & I & 0 \end{pmatrix} \begin{pmatrix} r_{t-1}^m \\ r_{t-1}^n \\ r_{t-2}^m \\ r_{t-2}^n \\ \ldots \\ r_{t-k-1}^m \\ r_{t-k-1}^n \end{pmatrix} + \begin{pmatrix} \eta_t \\ 0 \\ \ldots \\ 0 \end{pmatrix} \qquad (17),$$

or simply $x_t = \Theta x_{t-1} + \upsilon_t$. Note that $E_t(x_{t+k}) = \Theta^k x_t$, so that for $e_1 = (1,0,\ldots,0)'$. Also note that for any two interest rates such that $k = n/m$ is an integer the EH implies that

$$r_t^n = \frac{1}{k}\sum_{i=0}^{k-1} E_t\left(r_{t+mi}^m\right) \qquad (18).$$

The EH thus can be expressed alternatively as

$$e_2^{'}x_t = \frac{1}{k}\sum_{i=0}^{k-1} e_1^{'}\Theta^{mi}x_t \qquad (19),$$

where $e_2 = (0,1,0,\ldots,0)'$.

The constraints that satisfy the EH are given by

$$a_T(\theta) \equiv e_2^{'} - \frac{1}{k}\sum_{i=0}^{k-1} e_1^{'}\Theta^{mi} = 0 \qquad (20).$$

Since there is no simple form for the Jacobian of these constraints, the constraints are calculated numerically for the use in the iterative procedure described above.

The above analysis assumes that short and long rates are integrated of order zero, I(0). Campbell and Shiller (1987) address the concern that interest rates are not stationary. Their test is based on the fact that (2) can be rewritten as

$$S_t = E_t\sum_{i=1}^{k-1}(1 - i/k)\Delta^m r_{t+mi}^m \qquad (21),$$

where S_t is the spread between the long and the short rates, $r_t^n - r_t^m$, and Δ^m is an m-horizon change, that is, $\Delta^m w_t = w_{t+m} - w_t$. Campbell and Shiller propose to estimate a VAR of the form

$$x_t = A(L)x_{t-1} + w_t \qquad (22),$$

where $x_t = (\Delta r_t^m, S_t)'$ and A(L) is a P-order polynomial in the lag operator L, and to test the restrictions implied by (21). Equation (22) can be rewritten as

$$\begin{pmatrix} x_t \\ x_{t-1} \\ \dots \\ x_{t-P} \end{pmatrix} = \begin{pmatrix} A_1 & A_2 & \dots & A_P \\ I & 0 & \dots & 0 \\ 0 & I & \dots & \dots \\ \dots & \dots & \dots & \dots \\ 0 & \dots & I & 0 \end{pmatrix} \begin{pmatrix} x_{t-1} \\ x_{t-2} \\ \dots \\ x_{t-P-1} \end{pmatrix} + \begin{pmatrix} \omega_t \\ 0 \\ \dots \\ 0 \end{pmatrix} \qquad (23),$$

or more compactly as $x_t^* = Ax_{t-1}^* + \omega_t$ (24).

Campbell and Shiller (1987, 1991) note that (21) can be rewritten as

$$S_t = e_1^{'} A\left[I - (m/n)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1} x_t^* \qquad (25).$$

Therefore, the EH can be tested under the assumption that interest rates are not stationary by testing the following restriction:

$$e_{22}^{'} - e_1^{'} A\left[I - (m/n)(I - A^n)(I - A^m)^{-1}\right](I - A)^{-1} = 0 \qquad (26).$$

Campbell and Shiller (1987) note that (5) and (22) are comparable ($w_t = \eta_t$) only if the rates are cointegrated and the cointegrating vector is (1, -1). In other words, the rates should satisfy the necessary conditions for the EH to hold in the long run. Campbell and Shiller's specification allows one to test whether the EH holds at frequencies that are of interest to policymakers.

The starting point is the test of the restrictions implied by the EH without controlling for the quarter-end and year-end preferred habitat effects. As discussed above, a VAR in levels contains the assumption that the interest rates are stationary. Because, as statistical tests suggest, this is most likely not true for our data, VAR-in-differences is a much more suitable choice. That is, the set of dependent variables would consist of the first differences (i.e., daily changes) of the short and long rates. However, as Campbell and Shiller (1991) notice,

estimating a VAR in first differences $\left(\Delta r_t^m, \Delta r_t^n\right)$ would lead to a loss of information on the relative levels of the short and long rates. Because the EH relates the levels of the short and long rates, a test based on a non-level-preserving transformation of the rates cannot be considered a true test of the EH. A form of the VAR that does preserve the information on the relative levels of interest rates would include a first difference of the short rate and a spread between the rates as the two endogenous variables $\left(\Delta r_t^m, S_t\right)$. The VAR restrictions imply that information at time *t*, with the exception of the information contained in the spread, should not help predict future changes in the short-term rate. We use this specification which was proposed by Campbell and Shiller (1987) and was discussed above.[18]

Table 11 reports LM statistics for all 33 pairs of interest rates (one-three months, one-six months, and one-twelve months for each of 11 currencies). No exogenous variables, such as PH dummies, are included in the VAR at this point. The number of lags in the VAR was determined by the Akaike information criterion. The number of lags is in the parentheses next to the LM statistics.

The LM test rejects the EH for most interest rate pairs. The EH could not be rejected for any of short-long rate pairs for GBP and for three- and six-month long rates in AUD.[19] Our preliminary results suggest that, although preferred habitat effects influence the relationship between one-month rate and longer interest rates, they do not change the conclusion with respect to the EH. The output of Table 11 demonstrates that the EH does not hold for the

[18] VARs with daily changes in both rates were also run; the EH was always strongly rejected (results not reported).
[19] VARs in levels was run for CAD because Canadian LIBOR was found to be stationary, but these VARs are usually unstable. Thus, we report results for VAR-in-differences for CAD in Table 10. Given that it may be stationarity, the results for CAD are inconclusive when it comes to testing the EH.

majority of interest rate pairs when the one-month rate is the short rate, regardless of whether the currency exhibits a significant quarter-end/year-end effect in a one-month maturity. Of course, we have to be very careful making any conclusions before analyzing the effect of the PH variables on the LM statistics for the VARs.

Table 11. LM statistics for the EH tests, no PH controls.

Long rate / Currency	3m	6m	12m
USD			
- LM statistics	14.47 (2)	16.47 (2)	14.65 (2)
-p-value	0.006	0.002	0.006
GBP			
- LM statistics	11.42 (5)	9.57 (4)	9.35 (4)
-p-value	0.326	0.297	0.499
AUD			
- LM statistics	10.62 (2)	9.178 (4)	6.37 (3)
-p-value	0.031	0.328	0.383
Euro			
- LM statistics	30.37 (5)	46.74 (5)	26.38 (5)
-p-value	0.001	0.000	0.003
SF			
- LM statistics	31.91 (2)	32.19 (4)	N/A
-p-value	0.000	0.000	N/A
JPY			
- LM statistics	15.98 (2)	8.23 (2)	103.11 (2)
-p-value	0.003	0.084	0.000
CAD			
- LM statistics	11.24 (4)	19.25 (3)	17.74 (3)
-p-value	0.188	0.004	0.007
DM			
- LM statistics	27.31 (2)	15.37 (2)	17.69 (3)
-p-value	0.000	0.004	0.007
FF			
- LM statistics	26.46 (4)	31.83 (4)	457.62 (4)
-p-value	0.001	0.000	0.000
SP			
- LM statistics	23.08 (5)	26.68 (4)	250.38 (3)
-p-value	0.010	0.001	0.000
ITL			
- LM statistics	24.39 (2)	65.61 (5)	N/A
-p-value	0.000	0.000	N/A

Note: the number of lags in a VAR is in the parenthesis next to the coefficient

The set of the preferred habitat dummy variables used in the VAR is similar to that used in equation (1). When PH dummy variables are added to the VAR, turn-of-the quarter dummies are never significant. This is no surprise given that the turn-of-the-quarter effect in the one-month maturity was found to be small and often statistically insignificant. Given this, we only left the turn-of-the-year dummy variables (BYCR, AYCR, BYEND, and AYEND) in the estimation of the VARs. When PH dummy variables are added to the estimation of VAR, the procedure often does not converge, i.e., the restricted VAR cannot be estimated. It appears that the non-convergence becomes more likely as more variables are added. This is not related to the model stability: the largest roots of the unrestricted VARs do not tend to be larger than the roots of VARs without PH dummies.

The non-convergence is interpreted as a rejection of the EH. The iterative method used is a modification of the Newton method for numerical optimization. Newton's method is widely known to have good convergence properties when started near an optimum point. It is plausible to think that Bekaert and Hodrick's (2001) modification of it would also have good convergence properties near an optimum. Here we start with the unrestricted set of coefficients as the initial condition; it appears to be a natural choice. Thus, the non-convergence of the method is likely due to unrestricted VAR coefficients being "too far" from any set of coefficients that satisfy the EH. Hence, we believe that even if a different method were able to derive a set of restricted coefficients, their distance from the unrestricted set would most likely cause the Lagrange multiplier to reject the EH.

We next proceed to test the EH using the same procedure (the LM test based on GMM VAR estimations) for maturities with no preferred habitat effects, i.e., those that exceed one month. Since we use three-, six-, and twelve-month LIBOR in this study, three pairs of short-

long rates can be formed for each currency: three-six months, three-twelve months, and six-twelve months. These tests do not involve PH variables and will help us determine whether the EH holds for pairs of interest rates that involve only maturities longer than one month.

The output of the tests is reported in Table 12. It is very similar to the output of Table 11 in that the EH is rejected for the overwhelming majority of short-long interest rate pairs. Most of the rejections are significant at the 1% level, with a few more being significant at the 5% level. The EH is now rejected for GBP, although it held when the one-month rate was the short-term rate. The EH for CAD with three- and six-month rates can only be rejected at the 10% level. For one pair of interest rates - three-twelve months in CAD – the EH could not be rejected. Several times (three-six and six-twelve months for JPY, three-twelve and six-twelve months for FF and ITL) the unrestricted model could not be estimated due to non-convergence. Again, we interpret it as a strong rejection of the EH.

The output reported in Table 12 suggests that preferred habitat for liquidity in the one-month LIBOR prior to the year-end is unlikely to be responsible for the failure of the EH. LIBOR maturities involved in tests whose results are reported in Table 12 do not have quarter-end or year-end effects, and yet the EH is strongly rejected by the LM statistics for the overwhelming majority of interest rate pairs.

Table 12. LM statistics for the EH tests for LIBOR maturities exceeding one month.

Long rate	6m	12m	Long rate	6m	12m
Currency/ short rate			**Currency/ short rate**		
USD 3m			**CAD 3m**		
- LM statistics	27.98 (2)	25.93 (2)	- LM statistics	14.42 (4)	15.24 (5)
-p-value	0.000	0.000	-p-value	0.071	0.124
USD 6m			**CAD 6m**		
- LM statistics		47.23 (2)	- LM statistics		31.97 (5)
-p-value		0.000	-p-value		0.000
GBP 3m			***DM 3m***		
- LM statistics	22.00 (4)	19.94 (5)	*- LM statistics*	*60.63 (2)*	*20.40 (2)*
-p-value	0.005	0.030	*-p-value*	*0.000*	*0.000*
GBP 6m			***DM 6m***		
- LM statistics		21.04 (4)	*- LM statistics*		*56.87 (2)*
-p-value		0.007	*-p-value*		*0.000*
Euro 3m			**FF 3m**		
- LM statistics	*33.46 (5)*	*16.69 (4)*	- LM statistics	18.86 (5)	N/A
-p-value	*0.000*	*0.033*	-p-value	0.042	--
Euro 6m			**FF 6m**		
- LM statistics		*32.62 (5)*	- LM statistics		N/A
-p-value		*0.000*	-p-value		--
SF 3m			**SP 3m**		
- LM statistics	*40.81 (3)*	*16.39 (2)*	- LM statistics	21.60 (5)	172.12 (2)
-p-value	*0.000*	*0.003*	-p-value	0.017	0.000
SF 6m			**SP 6m**		
- LM statistics		*113.01 (2)*	- LM statistics		73.15 (5)
-p-value		0.000	-p-value		0.000
JPY 3m			**ITL 3m**		
- LM statistics	*N/A*	*18.54 (2)*	- LM statistics	29.47 (5)	N/A
-p-value	--	*0.005*	-p-value	0.001	--
JPY 6m			**ITL 6m**		
- LM statistics		*N/A*	- LM statistics		N/A
-p-value		--	-p-value		--
AUD 3m					
- LM statistics	26.03 (5)	18.38 (5)			
-p-value	0.004	0.049			
AUD 6m					
- LM statistics		40.98 (5)			
-p-value		0.000			

The number of lags in a VAR is in parenthesis next to the LM statistics.

2.3.4 Error-correction Models

If two variables are cointegrated, there is a long-term relationship that characterizes the equilibrium state. As we have seen, one-month rates are cointegrated with longer term interest rates within each currency. The normalized cointegrating vectors, although sometimes statistically different from the theoretical value consistent with the EH, (1, -1), are always close to this value. Moreover, the spreads between the one-month and longer-term LIBOR are stationary. While the EH appears to hold in the data in the long run, the deviations from the long-term equilibrium may persist for too long. It may be the reason why the EH has been so decisively rejected by the LM test.

Given a cointegrating vector normalized with respect to the short rate, (1, -β), the long-term equilibrium is attained when $r_t^m - \beta r_t^n = 0$, that is, when the short rate equals beta times the long-term rate. Short- and long-term rates fluctuate in response to stochastic shocks causing deviations from this equilibrium state. An error-correction model can be used to determine how and how quickly the long-term equilibrium is restored when deviations occur. For example, if the deviation is negative (i.e., $r_t^m - \beta r_t^n < 0$), the short-term rate must ultimately rise relative to the long-term rate. This gap may be closed in one of the following ways: (1) an increase in a short rate and/or a decrease in the long rate, (2) a rise in a long rate but a larger rise in a short rate, and (3) a fall in the short rate but a larger fall in a long rate. The error-correction model allows to determine the short-term dynamics for restoring the long-term equilibrium. If both interest rates are I(1), a simple error-correction model could be applied to the term structure:

$$\begin{aligned} \Delta r_t^m &= \alpha_m (r_{t-1}^m - \beta r_{t-1}^n) + \varepsilon_m \\ \Delta r_t^n &= \alpha_n (r_{t-1}^m - \beta r_{t-1}^n) + \varepsilon_n \end{aligned} \qquad (27).$$

The terms ε_m and ε_n are white-noise disturbances that may be correlated. The term in the parentheses represents the previous day's deviation from the long-term cointegrating relationship; over the long-term, it is zero. The two-variable error-correction model described by (23) is just a bivariate VAR in first differences augmented by the error-correction terms $\alpha_m\left(r_t^m - \beta r_t^n\right)$ and $\alpha_n\left(r_t^m - \beta r_t^n\right)$. The coefficients α_m and α_n measure the speed of adjustment of the short-term and long-term rate, respectively. The larger the absolute value of α_m (α_n), the greater the response of the short-term (long-term) rate to the previous day's deviation from the long-run equilibrium. Conversely, small absolute values of α_m (α_n) are a sign of unresponsiveness of the short-term (long-term) rate to deviations from equilibrium. Notice that the LHS of both equations in (27) is I(0) because the daily changes in interest rates (i.e., first differences) have been found to be stationary. In order for the error-correction model to be valid, the RHS must also be I(0). In other words, we reemphasize that there must be a cointegrating relationship between the short- and long-term rates in order to obtain sensible results. This is certainly true about the interest rates in our sample, as all of the interest rates series with the possible exception of CAD are integrated of order one, with stationary first differences and existing cointegrating relations between short- and long-term rates. The form (27) only has the error-correction term but it can and usually does include lagged differences. The model can also contain exogenous variables, for example, preferred habitat dummies. All regressors must be the same in both equations in order to obtain valid results.

Table 13 contains parameters α_m and α_n for each of the pairs of the short-term and long-term rates for each currency. At this point, no PH dummies are used in the model. The number of lags for error-correction models has been determined using the Akaike information criterion.

Table 13. Speed of adjustment parameters in error-correction models characterizing the relations between one-month LIBOR and longer-term LIBOR.

Long rate	3m		6m		12m	
Currency	α_m	α_n	α_m	α_n	α_m	α_n
USD	-0.066	-0.024	-0.033	-0.013	-0.015	-0.006
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.012)
GBP	-0.074	-0.043	-0.031	-0.018	-0.014	-0.008
	(0.000	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
AUD	-0.059	-0.036	-0.024	-0.015	-0.010	-0.006
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Euro	-0.046	-0.010	-0.025	-0.009	-0.014	-0.005
	(0.000)	(0.072)	(0.000)	(0.001)	(0.000)	(0.000)
SF	-0.043	-0.003	-0.020	0.000	-0.009	0.000
	(0.000)	(0.585)	(0.000)	(0.708)	(0.005)	(0.962)
JPY	-0.069	-0.029	-0.038	-0.017	-0.020	-0.008
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
CAD	-0.086	-0.048	-0.036	-0.018	-0.018	-0.009
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
DM	-0.036	-0.012	-0.017	-0.008	-0.009	-0.004
	(0.000)	(0.030)	(0.000)	(0.006)	(0.000)	(0.009)
FF	-0.090	-0.026	-0.045	-0.013	-0.025	-0.006
	(0.000)	(0.013)	(0.000)	(0.000)	(0.000)	(0.000)
SP	-0.368	-0.161	-0.190	-0.060	-0.118	-0.024
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
ITL	-0.086	-0.019	-0.056	-0.013	-0.025	-0.010
	(0.000)	(0.044)	(0.000)	(0.008)	(0.000)	(0.000)

The coefficient of the short-term rate has been set equal to one. P-values are in parentheses below coefficients.

As evident from Table 13, the short-term rate's response to the disequilibrium is in the expected direction. The negative sign of α_m implies that when $r_t^m - \beta r_t^n > 0$, the short rate typically decreases (that is, the dependent variable in the first equation of the VAR, the change in the short-term rate, is negative). Accordingly, the short rate increases when $r_t^m - \beta r_t^n < 0$. This is consistent with restoring the long-term equilibrium relationship. However, the long-term rate changes in the same direction as the short-term rate, as the signs of α_n's are also negative. This means that when the short rate increases to restore the equilibrium, the long rate also increases. The magnitude of changes of the long-term rate is always smaller, which means that the movement of the long-term rate only partially offsets

the movement by the short-term rate to restore the equilibrium. Thus, the long-term equilibrium is restored with an increase (decrease) in the short rate and a commensurate smaller increase (decrease) in the long rate. In the case of SF, the long-term rates do not respond to deviations from the equilibrium, only the one-month rate does.

Speed-of-adjustment parameters (the alphas) vary noticeably among currencies and maturities. For example, the largest α_m (in absolute value) is for the interest rate pair of one-month and three-month LIBOR for SP; it is equal to -0.376. That means that the one-month LIBOR for SP responds to the deviation from the long-term relationship by changing by more than one-third of that deviation the next day in the direction consistent with restoring the equilibrium. However, the three-month SP LIBOR change, expressed by α_n, is also quite large, a 15.4% of the deviation in the same direction as the short rate, offsetting the move of the short rate toward the equilibrium by almost a half. For other currencies, the speed-of-adjustment parameters are much smaller in absolute values, suggesting that the long-term equilibrium is restored much more slowly. For example, in the case of the one-three month rates for Euro, only 3.6% of the previous day's deviation from the equilibrium is eliminated the next day (the one-month LIBOR changes by 4.6% while the long-term rate changes by 1% in the same direction).

Another easily noticeable characteristic of the output from Table 13 is that the response of both short- and long-term interest rates to the disequilibrium lessens as the maturity gap between the two rates widens. For example, in the one-three month pair for JPY, the short rates' response is 6.9% of the deviation from equilibrium, and the three-month rate's response is 2.9% in the same direction. However, in the one-six months pair for JPY, the one-month LIBOR adjusts by only 2%, and the six-month rate changes by 0.8% of the

deviation in the same direction. This behavior is common for all of the currencies. It is likely to be a sign of cointegrating relationships weakening as the maturity of the long rate increases. It parallels the output of the augmented Dickey-Fuller test reported in Table 8.

Overall, the slow adjustments of interest rates back to the equilibrium may be the reason for rejection of the EH in our data. Although it is not clear what set of the speed-of-adjustment parameters would lead to non-rejection of the EH, adjustments of the magnitude of 3-6% per day may be interpreted as being too slow for the EH to hold when the maturity of the short-term rate is only one month.

As noted above, other (exogenous) variables may be added to the error-correction model. We proceed by incorporating turn-of-the-year preferred habitat dummies similar to those in equation (1) into the error-correction model. This will allow us to tell whether the speed of adjustment back to the equilibrium changes after accounting for the year-end effect in the one-month rate. Because the regressors in both equations must be the same, we insert the PH dummies into both equations in (27). These dummies are expected to have insignificant coefficients in the equation for the changes in long-term rate since three-, six-, and twelve-month rates do not have preferred habitat effects. The error-correction model will look as follows:

$$\begin{aligned} \Delta r_t^m &= \alpha_m (r_{t-1}^m - \beta r_t^n) + c_1 BYCR + c_2 AYCR + c_3 BYEND + c_4 AYEND + \varepsilon_m \\ \Delta r_t^n &= \alpha_n (r_{t-1}^m - \beta r_t^n) + c_1 BYCR + c_2 AYCR + c_3 BYEND + c_4 AYEND + \varepsilon_n \end{aligned} \qquad (27').$$

Panel A of Table 14 reports the speed-of-adjustment parameters along with the turn-of-the-year dummy coefficients from the error-correction model. The PH dummy coefficients contained in Panel B are from the first equation in (27') with the three-month LIBOR as the long-term rate. Of course, the coefficients of these variables for the one-month rate change

from the other two models (with six- and twelve-month LIBOR as long-term rates) are very close to those reported in the table.

Table 14. Speed of adjustment parameters in error-correction models for pairs of short-term (one-month) and long-term LIBOR (the turn-of-the-year dummies included).

Panel A. Speed-of-adjustment parameters

Long rate	3m		6m		12m	
Currency	α_m	α_n	α_m	α_n	α_m	α_n
USD	-0.052	-0.020	-0.030	-0.012	-0.015	-0.005
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.015)
GBP	-0.072	-0.041	-0.031	-0.017	-0.014	-0.007
	(0.000	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
AUD	-0.065	-0.039	-0.026	-0.015	-0.011	-0.006
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.002)
Euro	-0.046	-0.010	-0.026	-0.009	-0.014	-0.005
	(0.000)	(0.052)	(0.000)	(0.001)	(0.000)	(0.000)
SF	-0.026	-0.002	-0.014	0.000	-0.006	0.000
	(0.001)	(0.810)	(0.005)	(0.892)	(0.027)	(0.798)
JPY	-0.068	-0.029	-0.036	-0.016	-0.019	-0.006
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.006)
CAD	-0.083	-0.042	-0.035	-0.016	-0.016	-0.008
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
DM	-0.031	-0.008	-0.016	-0.007	-0.009	-0.004
	(0.000)	(0.143)	(0.000)	(0.017)	(0.000)	(0.028)
FF	-0.116	-0.048	-0.056	-0.019	-0.030	-0.008
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
SP	-0.320	-0.143	-0.153	-0.049	-0.094	-0.019
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
ITL	-0.094	-0.031	-0.059	-0.016	-0.027	-0.013
	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)

Panel B. Turn-of-the year dummy variable coefficients

	BYCR	AYCR	BYEND	AYEND
USD	0.258	0.014	-0.191	-0.036
	(0.000)	(0.196)	(0.000)	(0.001)
GBP	0.032	-0.013	-0.032	0.040
	(0.050)	(0.418)	(0.048)	(0.016)
AUD	0.023	0.009	0.001	-0.019
	(0.052)	(0.456)	(0.912)	(0.095)
Euro	0.111	0.028	-0.032	0.009
	(0.000)	(0.043)	(0.019)	(0.503)
SF	0.172	0.023	-0.117	-0.027
	(0.000)	(0.151)	(0.000)	(0.070)
JPY	0.048	0.009	-0.084	-0.013
	(0.000)	(0.327)	(0.000)	(0.180)
CAD	0.053	0.006	-0.036	-0.002
	(0.006)	(0.774)	(0.061)	(0.934)
DM	0.159	0.021	-0.094	-0.028
	(0.000)	(0.124)	(0.000)	(0.031)
FF	0.053	0.084	0.043	0.114
	(0.256)	(0.068)	(0.348)	(0.018)
SP	-0.029	0.025	-0.046	-0.011
	(0.851)	(0.870)	(0.762)	(0.943)
ITL	0.109	0.029	-0.173	-0.120
	(0.207)	(0.737)	(0.046)	(0.191)

The coefficient of the short-term rate has been set equal to one. P-values are in parentheses below coefficients.

Results reported in Table 14 suggest that the addition of the turn-of-the-year dummies almost does not change the speed-of-adjustment parameters. The coefficients of variables BYCR and BYEND are significant for USD, Euro, SF, JPY, and DM, as expected. They are consistent with the year-end preferred habitat for liquidity in these currencies. They are also statistically significant (although much smaller in magnitude) for GBP and (marginally) for CAD. These coefficients are interpreted as changes in interest rates (e.g., 0.26 means a 26 basis point increase in the one-month USD LIBOR).[20]

[20] One can argue that the PH dummies used in (23') do not cover the entire period of increased short-term rates. To address this concern, we conducted the same tests using YPH (which covers the entire period of increased one-month rates prior to the year-end) instead of these dummies. The results are qualitatively similar to those reported in Table 14.

2.4 Conclusion

Empirical studies suggest that the expectations hypothesis of the term structure of interest rates is often rejected, especially at the short end. There are many possible reasons for this rejection: time-varying risk premia, irrationality of market participants causing failure of rational expectations, overreaction of market participants to monetary policy changes. In the light of our earlier findings, preferred habitat for liquidity in the short-term rate may also contribute to the failure of the EH at the short end of the term structure. However, our results suggest that, although the year-end preferred habitat for liquidity certainly has influence on short-term (one-month) interest rates in major world currencies, it is not responsible for the rejection of the expectations hypothesis for the major world currencies.

The one-month LIBOR is cointegrated with longer-term (three-, six-, and twelve-month) LIBOR maturities within each currency. It implies the existence of a long-run equilibrium between short- and long-term interest rates. However, the interest rates appear to move too slowly to restore the long-run equilibrium. Short-term and long-term rates tend to move in the same direction, with the magnitude of the short-term rate movements exceeding that of the long-term rate movements. This means that when the short-term rate moves to restore the long-term equilibrium, the long-term rate tends to partially offset it by moving in the same direction. This slow restoration of the long-term equilibrium may cause the failure of the Expectations Hypothesis in the short run.

REFERENCES

Allen, A., and A. Saunders, 1992, Bank window dressing: Theory and evidence, *Journal of Banking and Finance* 16, 583-623.

Ariel, R., 1987, A monthly effect in stock returns, *Journal of Financial Economics* 18, 161-174.

Banz, R., 1981, The relationship between returns and market value of common stocks, *Journal of Financial Economics*, 9, 3-18.

Bekaert, G., R. Hodrick, and D. Marshall, On biases in tests of the expectations hypothesis of the term structure of interest rates, *Journal of Financial Economics* 44, 309-348.

Blume, M., and I. Friend, 1986, Recent and prospective trends in institutional ownership and trading of exchanges and OTC stocks, *unpublished University of Pennsylvania working paper*.

Brown, R., K. Cyree, M. Griffiths, and D. Winters, 2004, Further analysis of the expectations hypothesis using very short-term rates, *Texas Tech University working paper*.

Campbell, J., 1986, A Defence of traditional hypotheses about the term structure of interest rates, *Journal of Finance* 41, 183-193.

Campbell, J., and R. Shiller, 1991, Yield spreads and interest rate movements: A bird's eye view, *Review of Economic Studies* 58, 495-514.

Chan, K. C. and N. Chen, 1988, An unconditional asset-pricing test and the role of firm size as an instrumental variable for risk, *Journal of Finance*, 43, 309-326.

Cook, T., and T. Hahn, 1990, Interest rate expectations and the slope of the money market yield curve, *Federal Reserve Bank of Richmond Economic Review* 76 (5), 3-26.

Cox, J., J. Ingersoll, and S. Ross, 1981, A re-examination of traditional hypotheses about the term structure of interest rates, *Journal of Finance* 36, 769-799.

Cox, J., J. Ingersoll, and S. Ross, 1985, A theory of the term structure of interest rates, *Econometrica* 53, 385-408.

Cuthbertson, K., 1996, The expectations hypothesis of the term structure: the UK interbank market, *The Economic Journal* 106, 578-592.

Cyree, K. and D. Winters, 2001, Analysis of the federal funds rate changes and variance patterns, *Journal of Financial Research* 24, 403-418.

Dimson, E., 1979, Risk mearurement when shares are subject to infrequent trading, *Journal of Financial Economics* 7, 197-216.

Downing, C., and S. Oliner, 2004, The term structure of commercial paper rates, *Federal Reserve Board* working paper.

Enders, W., 1995, *Applied Econometric Time Series*, John Wiley and sons.

Fisher, I., 1896, Appreciation and interest, *Publications of the American Economic Association* 11, 21-29.

Fisher, M., and C. Gilles, 1998, Around and around: The expectations hypothesis, *Journal of Finance* 53, 365-383.

Griffiths, M. and R. White, 1993, Tax induced trading and the turn-of-the-year anomaly: An intraday study, *Journal of Finance* 48, 575-598.

Griffiths, M. and D. Winters, 2004, The-turn-of-the-year in the money markets: Tests of the risk-shifting window dressing and preferred habitat hypotheses, *Journal of Business*, forthcoming.

Griffiths, M. and D. Winters, 2003, The year-end price of risk in a market for liquidity, *The American Graduate School of International Management working paper*.

Griffiths, M. and D. Winters, 1997, On a preferred habitat for liquidity at the turn-of-the-year: Evidence from the term-repo market, *Journal of Finance Services Research* 12, 21-38.

Griffiths, M., and D. Winters, 1995. Day-of-the-week effects in federal funds rates: Further empirical findings, *Journal of Banking and Finance* 19, 1265-1284.

Hall, A., H. Anderson, and C. Granger, 1992, A cointegration analysis of Treasury bill yields, *Review of Economics and Statistics*, 74, 116-126.

Hamilton, J., 1996, The daily market for federal funds, *Journal of Political Economy* 104, 26-56.

Hansen, L., 1982, Large sample properties of generalized method of moments estimators, *Econometrica* 50, 1029-1054.

Hardouvelis, G., 1994, The term structure spread and future changes in long and short rates in G7 countries: Is there a puzzle? *Journal of Monetary Economics* 33, 255-283.

Hardouvelis, G., 1988, The predictive power of the term structure during recent monetary regimes, *Journal of Finance* 43, 339-356.

Haugen, R. and J. Lakonishok, 1988, The incredible January effect, Homewood, IL: Dow Jones-Irwin.

Hicks, J., 1946, Value and capital, 2nd edition, Oxford University Press.

Hurn, S., T. Moody, and A. Muscatelli, 1995, The term structure of interest rates in the London interbank market, *Oxford Economic Papers* 47, 418-436.

Johansen, S. 1991, Estimation and hypothesis testing of cointegration vectors in gaussian vector autoregressive models, *Econometrica* 59, 1551–1580.

Johansen S. 1995, Likelihood-based Inference in cointegrated vector autoregressive models, Oxford University Press.

Jones, S., W. Lee, and R. Apenbrink, 1991, New evidence on the January effect before personal income taxes, *Journal of Finance* 46, 1909-1924.

Jordan, S. and B. Jordan, 1991, Seasonality in daily bond returns, *Journal of Financial and Quantitative Analysis*, 26, 269-285.

Keim, D., 1983, Size-related anomalies and stock return seasonality: Further empirical evidence, *Journal of Financial Economics*, 12, 13-32.

Lakonishok, J., A. Shleifer, R. Thaler, and R. Vishny, 1991, Window dressing by pension fund managers, *American Economic Review* 81, 227-231.

Lakonishok, J., and S. Smidt, 1984, Volume and turn of the year behavior*, Journal of Financial Economics* 13, 435-456.

Lakonishok, J., and S. Smidt, 1988, Are seasonal anomalies real? A ninety-year perspective, *Review of Financial Studies* 1, 403-425.

Lange, J., B. Sack, and W. Whitesell, 2003, Anticipation of monetary policy in financial markets, *Journal of Money, Credit and Banking* 35, 889-909.

Longstaff, F., 2000a, Arbitrage and the expectations hypothesis, *Journal of Finance* 55, 989-994.

Longstaff, F., 2000b, The term structure of very short-term rates: New evidence for the expectations hypothesis, *Journal of Financial Economics* 58, 397-415.

Modigliani, F., and R. Sutch, 1966, Innovations in interest rate policy, *American Economic Review* 56, 178-197.

MucCulloch, J., 1993, A reexamination of traditional hypothesis about the term structure: A comment, *Journal of Finance* 48, 779-789.

Musto, D., 1997, Portfolio disclosures and year-end price shifts, *Journal of Finance* 52, 1563-1588.

Musto, D., 1999, Investment decisions depend on portfolio disclosures, *Journal of Finance* 54, 935-952.

Ogden, J., 1987, The end of the month as a preferred habitat: a test of operational efficiency in the money market, *Journal of Financial and Quantitative Economics* 22, 329-344.

Ogden, J., 1990, Turn-of-the-month evaluations of liquid profit and stock returns: A common explanation for the monthly and January effect, *Journal of Finance* 45, 1259-1272.

Park, S. and M. Reinganum, 1986, The puzzling price behavior of Treasury bills that mature at the turn of calendar months, *Journal of Financial Economics* 16, 267-283.

Patel, A., and G. Shouesmith, 2004, Term structure linkages surrounding the Plaza and Louvre accords: Evidence from Euro-rates and long-memory components, *Journal of Banking and Finance* 28, 2051-2075.

Reinganum, M., 1981, Misspecification of capital asset pricing: Empirical anomalies based on earnings yields and market values, *Journal of Financial Economics* 9, 19-46.

Roberts, W., D. Runkle, and C. Whiteman, 1996, A daily view of yield spreads and short-term interest rate movements, *Journal of Money, Credit and Banking* 28, 34-53.

Roll, R., 1981, A possible explanation of the small firm effect, *Journal of Finance* 36, 879-888.

Roll, R., 1983, Vas ist das? The turn-of-the-year effect and the return premia of small firms, *Journal of Portfolio Management* 9, 18-28.

Scholes, M., and J. Williams, 1977, Estimating betas from nonsynchronous data, Journal of Financial Economics 5, 309-327.

Simon, D., 1994, Further evidence on segmentation in the Treasury bill market, *Journal of Banking and Finance* 18, 139-151.

Simon, D., 1991, Segmentation in the Treasury Bill Market: Evidence from Cash Management Bills, *Journal of Financial and Quantitative Analysis* 26, 97-108.

Simon, D., 1990, *Expectations and the Treasury bill – federal funds spread over recent monetary policy regimes*, Journal of Finance 45, 567-577.

Stigum, M., 1990, *The Money Market*, Third Edition, Dow Jones-Irwin.

Thornton, D., 2004a, Tests of the expectations hypothesis: Resolving the anomalies when the short-term rate is the federal funds rate, *Federal Reserve Bank of St. Louis working paper*.

Thornton, D., 2004b, Tests of the expectations hypothesis: Resolving the Campbell-Shiller paradox, *Federal Reserve Bank of St. Louis working paper*.

Thornton, D., 2004c, Testing the expectations hypothesis: Some new evidence from Japan, *Federal Reserve Bank of St. Louis Review* 86 (5), 21-39.

Tzavalis, E., and M. Wickens, 1997, Explaining the failure of the term spread models of the rational expectations hypothesis of the term structure, *Journal of Money, Credit and Banking* 29, 364-80.