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The Impact of Chinese Purchases of U.S. Government Debt on the Treasury Yield Curve

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The Impact of Chinese Purchases of U.S. Government Debt on the Treasury Yield Curve

Radhames A. Lizardo and Andre Varella Mollick

Abstract

Examining monthly data from May of 1985 to May of 2008, we find that increases in Chinese purchases of U.S government debt lead to decreases in Treasury yields. The effect is stronger as the maturity increases: a one percent increase in purchases of U.S. Treasuries by Chinese investors lowers the two-year (ten-year) Treasury yield by 10 to 38 basis points (39 to 55 basis points) on average, *ceteris paribus* . Overall, the demand-side variable capturing Chinese purchases of U.S. Treasuries improves the cointegrating properties of U.S. interest rates. In-sample and out-of-sample forecasts reinforce that the model with Chinese purchases greatly outperforms basic models of the yield curve. This study has implications for the business world since we document that Chinese investors contribute to lower U.S. Treasury yields and thus to lower U.S. interest rates in general.

KEYWORDS: Chinese investors, interest rates, US Treasury yields, yield curve

1. Introduction

China is an example of an economy which has turned out to be a new engine of global or regional growth. Figure 1 shows the growth path of the Chinese economy relative to that of the United States. China has been growing at a ratio between three and four-to-one relative to the U.S. After 2001, China's economy began to diverge from that of the U.S. and kept moving upward while that of the U.S. moved below its historic 3% growth rate. Figure 2 documents an increasing Chinese current account surplus while the U.S. current account deficit has intensified. The current account surplus of China has climbed to over 9% of their GDP while the current account deficit of the U.S. has reached about 6%. The increasing competitiveness of China exports has elicited an ever increasing flow of foreign direct investment. As a result, China's total reserves have surpassed \$2.1 trillion in mid-2009.

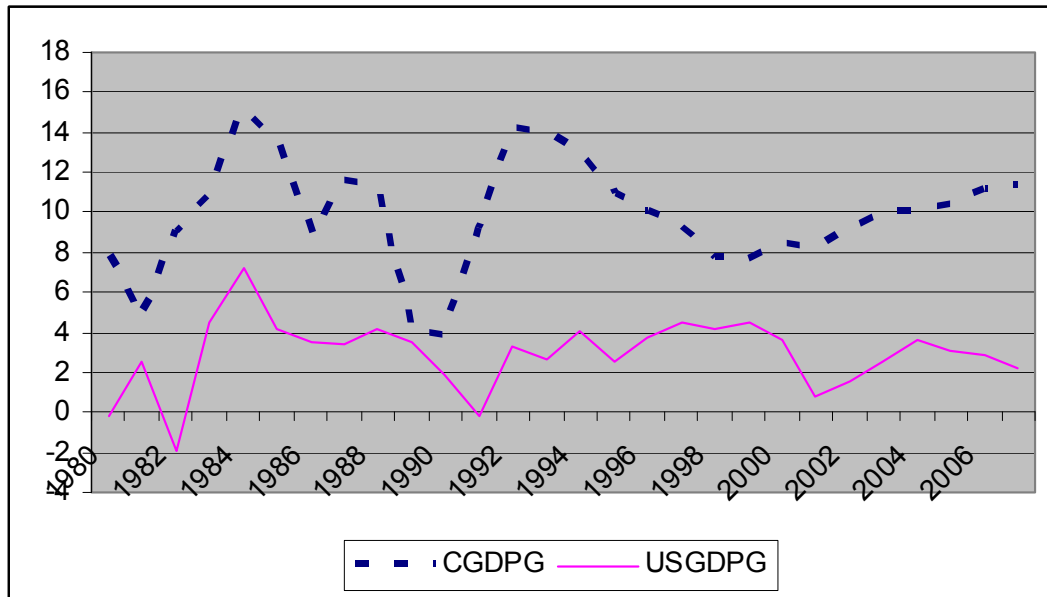
Friedman (2005) suggests that globalization has been leveling the competitive playing fields between industrial and emerging markets and that the convergence of technology is causing a shift in the pattern of global output. Such a process has been lifting countries from being insignificant players in world affairs to crucial participants in global finance. With these changes, the direction of the waves created by country-specific financial shocks is shifting. While the basic underpinnings of a "flat world" were originally envisaged within the context of trade and labor movements, there is nothing that precludes the concept to be applied in the financial world as well. In fact, since financial markets adjust faster to new information, the adjustment should in principle be more visible in finance than in the goods or labor market. For an extensive and economic-based review of the "world is flat" metaphor, see Leamer (2007). The redirection of financial flows from China into the developed world make us analyze in this paper the linkage between the Chinese purchases of U.S. Government Debt and the Treasury yield curve. For an example on the impact of China on labor market of Mexican maquiladoras, see Mollick and Wvally-Vázquez (2006); for impact of China on Latin American exports, FDI inflows, and terms of trade, see Jenkins et al. (2008).

The "flatter world" has made China recycle its trade surplus with the U.S. back into dollars, especially into U.S. Treasury bonds. China has become the largest financier of the U.S. huge budget and trade deficits. By mid-2009, China holding of U.S. Treasury Securities was approaching USD 800 billion (<http://www.ustreas.gov/tic/mfh.txt>), which represents close to 23% of the \$3,382 billion of U.S. Treasury Securities held by foreigners. China's share of U.S. Treasury Securities held by foreigners has been consistently increasing from close to 0% in 1985 to about 20% in June of 2008 and to around 23% by June of 2009.

China can exert a great deal of influence on both the value of the U.S. dollar and also on the U.S. Treasury yields. One question posed has been what if

China decides to get rid of the huge amount of U.S dollar assets it holds, including the possible negative effects on the U.S. economy.

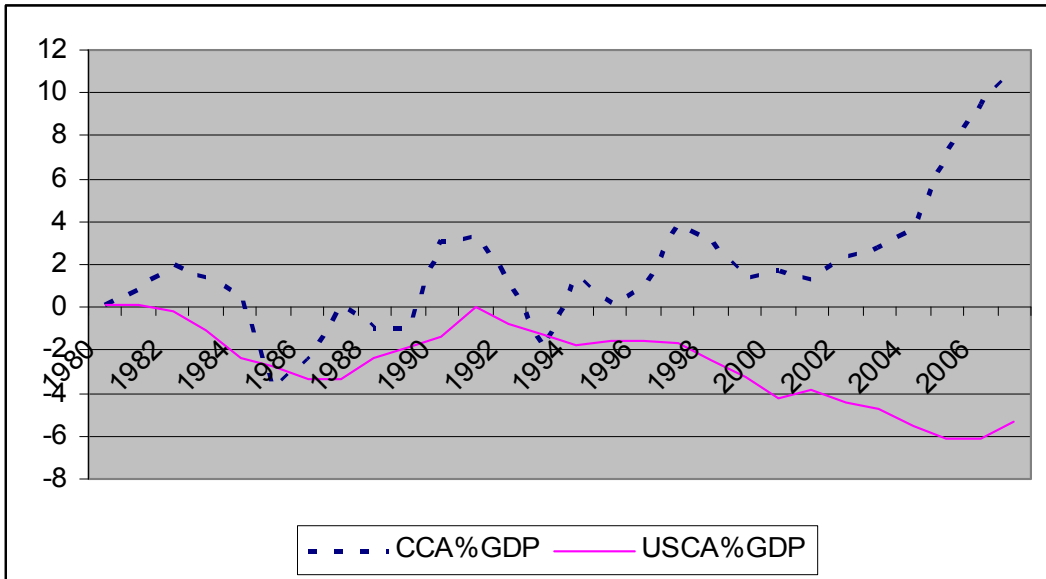
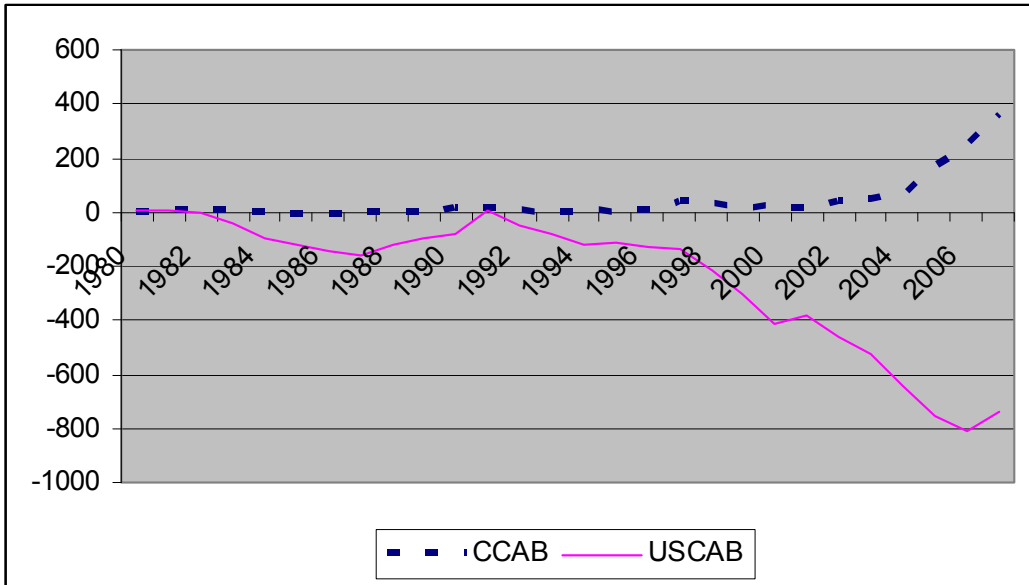
Figure 1. China's Gross Domestic Product Growth (CGDPG) vs. Gross Domestic Product Growth of the U.S. (USGDPG).



Notes: Constructed by the author, using data from International Financial Statistics (IFS) of the International Monetary Fund (IMF), downloaded from <http://www.imfstatistics.org>.

Figure 3 depicts the long-term trends between Purchases of U.S Treasuries by Chinese Investors (PUSTC) and the U.S 10-year Treasury Constant Maturity Rates (i10). An inverse relationship between these two series is observed: as PUSTC goes up, i10 goes down. It suggests that the amount of purchases of U.S. Treasuries by China is significant enough to put upward pressure on the price of the U.S 10-year Treasury security. As a result of the increased price, the yield (i10) goes down. Figure 4 shows rolling correlations between i10 and PUSTC based on a moving-window of 120-month periods. There is an increasing negative rolling correlation between PUSTC and i10. From the end of 2002 onwards the rolling correlation coefficients between these two variables have been consistently below -0.50.

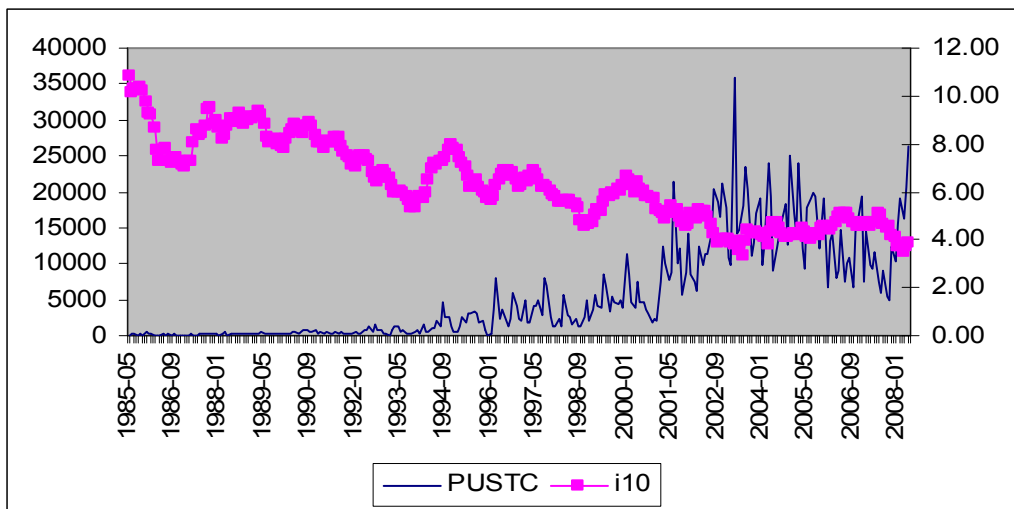
Figure 2. Current Account Balance for China (CCAB) and the United States (USCAB) and as Relative % of GDP.



Notes: Constructed by the author, using data from International Financial Statistics (IFS) of the International Monetary Fund (IMF), downloaded from <http://www.imfstatistics.org>.

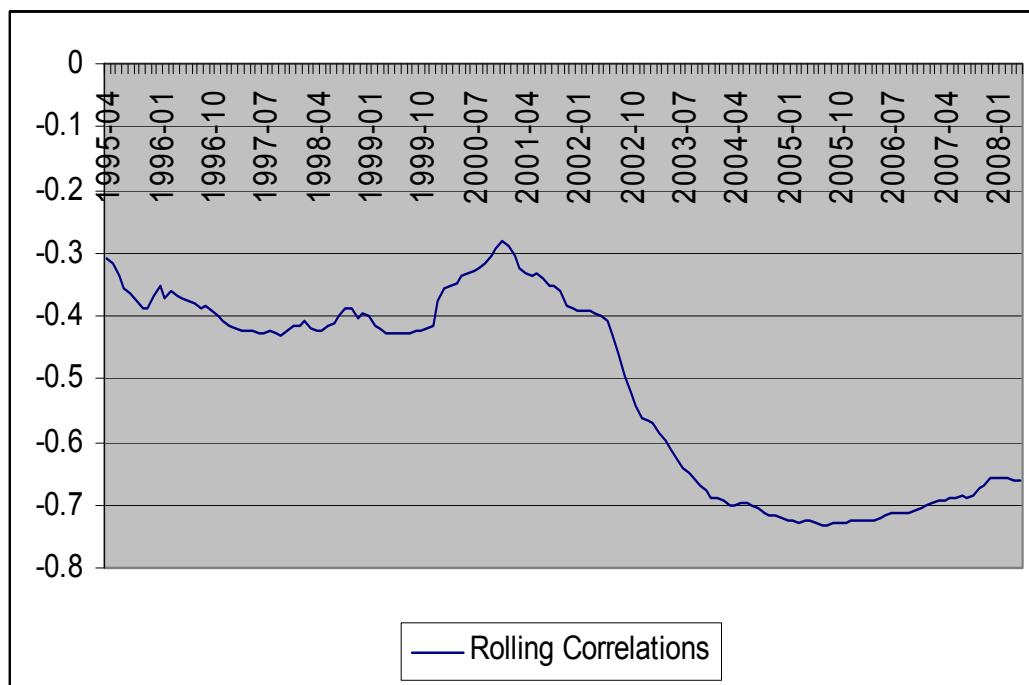
Contagion of regional and country specific financial shocks has been the subject of many studies. See, for example, Levin, (1974), Daniel (1981), Calomiris and Schweikart (1991), Chan et al. (1992), Karolyi (1995), Ammer and Mei (1996), and Baig and Goldfajn (1999). Because the U.S. economy has been a dominant global force since World War II, many of these studies have concentrated on analyzing transmission of U.S. economic shocks to other countries or regions; e.g., Eun and Shim (1989) and Cochran and Mansur (1991). Studies about how foreign economic shocks affect the U.S. are less common, perhaps because it has been broadly assumed that the U.S. economy is resistant to foreign shocks. During the 1990s, the U.S. economy was so strong and stable that many believed the U.S. economy had reached a “new paradigm”, as reviewed by Zuckerman (1998) and Krugman (2000). On the other hand, few studies explore spillover effects to the U.S. financial markets from foreign economic shocks. Peek and Rosengren (1997) find that the decline in Japanese stock prices resulted in a decrease in lending by Japanese banks in the U.S. Mollick and Soydemir (2008) find that a one-time increase in net Japanese purchases of U.S. Treasury securities has an immediate negative effect on U.S. long bond yields but short-lived yen depreciation. More recently, Hurley (2010) focusing on the holdings by a group of five Asian nations, found the presence of a long-run causal relations between the five Asia nations purchases of U.S. treasuries and the federal funds rate.

Figure 3. Long-Term Trends between Purchases by Chinese Investors of U.S. Treasuries (PUSTC) and i10 (left scale in USD millions, right scale in %).



Notes: Constructed by the authors, using data from the Federal Reserve Bank of St. Louis, downloaded from <http://www.frbstlouis.com> and Treasury International Capital (TIC) downloaded from <http://www.treas.gov/tic/>

Figure 4. Purchases of U.S. Treasuries and i10 Correlations using a Moving Window of 120-month Periods.



Note. Constructed by the authors, using data from the Federal Reserve Bank of St. Louis, downloaded from <http://www.frbstlouis.com> and *Treasury International Capital* (TIC) downloaded from <http://www.treas.gov/tic/>.

The relationship between the Chinese purchases and U.S. treasury yields is one that has been questioned a great deal by both academics and non-academics in recent years given the precipitous emergence of China as a potential threat to U.S. economic supremacy, yet, there has been little or no academic research done on the topic to date. This paper fills the gap. This study focuses on the impact that purchases of U.S Treasuries by China has on the U.S. Treasury Yield Curve. The United States has been enjoying a declining cost of borrowed funds. This is so because interest rates tend to co-move as documented by Sarno and Thornton (2003). Therefore, interests on corporate and other types of bonds have also moved downward. While the U.S. disinflation and U.S. monetary policy perhaps explain most of the downturn, it is also possible that exogenous forces (such as the recent Chinese appetite for U.S. fixed income assets) have had an impact. Cost of funds may be affected by other forces as well. Gompers et al. (2003), for example, develop a corporate governance index score based on several corporate governance provisions, with six of them serving as likely indicators of shareholder rights strength.

More recently, Collins and Huang (2011) find positive effects of management entrenchment on the cost of equity capital. We provide evidence in this paper - both with cointegration analysis and with in-sample and out-of-sample forecasts - that the purchases of U.S. Treasuries by Chinese have significantly lowered and flattened the U.S. Treasury Yield Curve. This study has implications for business world since we document that Chinese investors contribute to lower U.S. Treasury yields and thus to lower U.S. interest rates in general.

2. The Theoretical Framework and Methodology

Finance theory posits that interest rate on a debt security, such as a corporate bond, is determined by a real risk-free rate of interest, θ , plus several premiums that reflect expected inflation (EI), maturity risk premium (MRP), default risk premium (DRP) and liquidity risk premium (LRP). Following Brigham and Ehrhardt (2005), the interest rate function can be expressed as follows:

$$i_t = \theta_t + EI_t + MRP_t + DRP_t + LRP_t. \quad (1)$$

Because Treasuries have essentially no default or liquidity risk, we can impose $DRP = LRP = 0$ in (1). However, even Treasury bonds are exposed to a significant risk of price declines and a maturity risk premium is included in the yield function to reflect this risk on long-term securities. Therefore, the interest on a Treasury bond that matures in t years can be expressed as a function of the following determinants:

$$i_t = \theta_t + EI_t + MRP_t \quad (2),$$

where θ_t represents the interest rate that would exist on a riskless security if zero inflation were expected. Given the importance of expected inflation to the valuation of Treasuries and the fact that short-term rates highly co-move, we proxy θ_t with the effective federal funds rate. This formulation allows us to use the theoretical component of expected inflation for securities maturing in one year or longer. Maturity Risk Premium is derived as follows:

$$MRP_t = MAX[0, i_{LT} - i_{ST}] \quad (3),$$

where: i_{LT} represents the yield on the 10-year Constant Maturity Treasury yield and i_{ST} represent the yield on the 3-month Treasury bill rate. This term is present extensively in studies of the yield curve, such as Diebold and Li (2006).

Most of past empirical work focused on testing the Expectation Hypothesis (EH) of the term structure of interest rates using cointegration and equilibrium correction models. See, for example, Engle and Granger (1987), Simon (1990), Campbell and Shiller (1991), Hall et al. (1992), Engsted and Tanggaard (1994), Roberds et al. (1996), Lanne (2000), and Thornton (2005). The EH posits that the interest rates on long-term securities are simply a weighted average of current and expected future short-term interest rates. Several papers have explored the behavior of the shortest term rate (FF) and the 3-month TB and conclude that they behave in accordance with the EH. See, for example, Cook and Hahn (1989), Goodfriend (1991), Rudebush (2002), and Sarno and Thornton (2003).

This study focus on the impact that purchases of U.S Treasuries by China (PUSTC) has on the U.S. Treasury Yield Curve. We first estimate *the basic model* (2) as follows:

$$i_T = \beta_0 + \beta_1\theta_t + \beta_2EI_t + \beta_3MRP_t + \varepsilon_t \quad (4),$$

where: T = the One-year; Two-year; Three-year; Five-year; Seven-year; and Ten-year constant maturity yield, respectively. In symbols, i_1 , i_2 , i_3 , i_5 , i_7 , and i_{10} . We expand (4) and include the log of purchases of U.S Treasuries by China (PUSTC) as a predictor of U.S Treasury yields, in the *composite model* augmented by this factor:

$$i_T = \beta_0 + \beta_1\theta_t + \beta_2EI_t + \beta_3MRP_t + \beta_4 \log(PUTC_t) + \varepsilon_t \quad (5)$$

Theory suggests an inverse relationship between the interest rate and the price of a bond. Since interest rates co-move, when the very short-term rates move higher all other rates should follow to some degree and the price of existing bonds, *ceteris paribus*, should decrease (which results in an increased yield). Therefore we expect β_1 to be positive. The same applies to inflation premium (EI), and maturity risk premium (MRP) and we expect β_2 and β_3 to be positive as well. On the other hand, a *significant* increase in purchases of U.S. Treasuries by China should put upward pressure on the price of Treasuries, which should result in a lowered yield. As a result, we expect β_4 to be negative.

The empirical versions of (4) and (5) are estimated by OLS; dynamic OLS (DOLS) by Stock and Watson (1993); and by the multivariate maximum likelihood procedure of Johansen (1998, 1991): JOH-ML, which may be done using a Vector Error Correction Model (VECM) in the following form:

$$\Delta Y_t = \Pi Y_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \Phi D_t + \varepsilon_t, t = 1, \dots, T \quad (6)$$

If the long-run impact matrix Π in (6) is less than full rank, it can be decomposed as:

$$\Pi = \alpha\beta' \quad (7),$$

where α is an $n \times r$ matrix of speed of adjustments and β' is an $r \times n$ matrix of cointegrating coefficients. Block exogeneity tests are ruled out if the nonstationary variables are cointegrated (Enders 2004). The long-run impact matrix Π in (6) is less than full rank, it can be decomposed as in (7). If the speed of adjustments (α) is negative and statistically significant, one would conclude that the direction of causality goes from the regressors to the dependent variable.

The final step involves the comparison of the forecasting performance of the basic against the composite model. We use a recursive window to generate a series of out-of-sample forecasts for the last twelve months, the holdout sample. In a recursive forecasting model, the initial estimation date is fixed, but additional observations are added one at a time to the estimation period. The Diebold and Mariano (1995) statistic (DM) is reported, obtained by regressing the loss differential series on an intercept and a MA (1) term to correct for serial correlation. The mean square errors (MSE) are calculated as follows:

$$MSE = \frac{1}{T - (T_1 - 1)} \sum_{t=T_1}^T (y_{t+s} - f_{t,s})^2 \quad (8)$$

where T is the total sample size (in-sample + out-of-sample), and T_1 is the first out-of-sample forecast observation. In-sample model estimation initially runs from observation 1 to $(T_1 - 1)$, and observations T_1 to T are available for out-of-sample estimation, i.e. a total holdout sample of $T - (T_1 - 1)$. In addition, we calculate the Theil's (1966) U -statistic which is defined as follows:

$$U = \sum_{t=T_1}^T \sqrt{\frac{\left(\frac{y_{t+s} - f_{t,s}}{x_{t+s}}\right)^2}{\left(\frac{y_{t+s} - fb_{t,s}}{x_{t+s}}\right)^2}} \quad (9),$$

where: $fb_{t,s}$ is the forecast obtained from a benchmark model (the composite model in our analysis). A U -statistic of one implies that the model under consideration and the benchmark model have equal forecasting abilities while a

value of more than one implies that the benchmark model is superior to the basic model, and vice versa.

3. The Data and Descriptive Statistics

During the peak of the financial crisis in 2008, The Federal Reserve expanded the money supply by adding new assets and liabilities without sterilization. This “quantitative easing” contributed to the rising of U.S treasuries’ which lowered their yields. For that reason, this paper excludes data observations after May 2008 which as outliers could bias the results.

The data are monthly observations on the U.S 10-year, U.S 7-year, U.S 5-year, U.S 3-year, U.S 2-year and U.S 1-year Treasury Constant Maturity Rate yields (series ID: GS 10, GS 7, GS 5, GS 3, GS 2, and GS 1) from May, 1985 to May, 2008 which come from the Board of Governor of the Federal Reserve System, downloaded from the U.S. Federal Reserve of Saint Louis (<http://research.stlouisfed.org/fred2/categories/115>). The release is the H.15 “Selected Interest Rates”, monthly rate, in percentage and average of business days.

Monthly observations on the U.S effective Federal Funds (FF), which is a weighted average of the rates on federal funds transactions of a group of federal funds brokers who report their transactions daily to the Federal Reserve Bank of New York. The Series ID is the FEDFUNDS from May, 1985 to May, 2008 which come from the Board of Governor of the Federal Reserve System. The release is H.15 “Selected Interest Rates”, monthly rate, in percent and average of daily figures. Monthly calculations of the Maturity Risk Premium, are computed by subtracting the Series GS 10 from the Series TB3MS (3-month Treasury bill rate) for the months of 1985:5 to 2008:5, downloaded from the U.S. Fed Res of St. Louis from: (<http://research.stlouisfed.org/fred2/categories/118>); (<http://research.stlouisfed.org/fred2/categories/119>).

Monthly observations of Purchases of U.S. Treasury bonds by China (PUSTC) for the months from 1985:5 to 2008:5 [PubMed](#) come from the *Treasury International Capital* (TIC) and are downloaded from <http://www.treas.gov/tic/>. The TIC data represent foreign investor’s purchases and sales of U.S.’s long-term securities as reported by commercial banks, bank holding companies, brokers and dealers, foreign banks, and non-banking enterprises in the U.S.

Monthly observations of University of Michigan inflation expectation over the period from May of 1985 to May of 2008 come from the Board of Governor of the Federal Reserve System. The Series ID is MICH and was downloaded from <http://research.stlouisfed.or/fred2/categories/98>. In this measure of inflation expectations participants are asked what they expect inflation to be over the next 5 to 10 years. Market-based estimates of expected inflation based on the

difference between the nominal treasury notes and TIPS are available only starting in 1997. For details, see <http://www.clevelandfed.org/research/data/TIPS/bg.cfm>

4. Empirical Results

Since some unit root tests are more robust than others, we used the augmented Dickey and Fuller (1979) test, in addition to the modified ADF test proposed by Elliott et al. (1996), and the KPSS method suggested by Kwiatkowski et al. (1992). *FF*, *EI*, *MRP*, and *PUSTC* are clearly non-stationary series in levels. On the other hand, all series included are clearly stationary when first differenced. All series appear to be I (1): they have a unit root in levels, but are stationary when first differenced. Detailed unit root test results are available upon request.

The upper panel of Table 1 shows the cointegration test results of the basic model. There is strong support for the existence of a stable long-run relationship among i_t , θ_t , EI_t , and MRP_t as given by the Johansen (1988, 1991) trace and maximum eigenvalue tests. However, it is less clear the number of cointegration relationships, especially at the yields with shorter maturities. For instance, from one to three years there appears to be more than one cointegrating vector. The lower panel of Table 1 shows the cointegration test results of the composite model represented in (5). There is again strong support for the existence of a stable long-run relationship among i_t , θ_t , EI_t , MRP_t and $PUSTC_t$. However, only for one-year maturity (and for the trace test only) there is rejection of the null of “at most 1” cointegrating vector. Thus, for the composite model the data are perfectly consistent with a unique cointegration vector.

Table 1 suggests empirical support for both (4) and (5) as cointegrating relationships. However, we know theoretically that both of these equations cannot be cointegrating relationships. If $\log(PUSTC)$ is I (1) as concluded from the unit root tests and (5) is a cointegrating relationship, then (4) cannot also be a cointegrating relationship. The reason is that the disturbance term in (4) will be I (1), since $\log(PUSTC)$ becomes part of that disturbance term. But (4) cannot represent a cointegrating relationship with non-stationary disturbances: a contradiction. There is clearly an inconsistency in (4). Allowing the $PUSTC$ to be part of the cointegrating vector one can obtain only one cointegrating vector as depicted in (5) and supported by the data in the lower panel of Table 1.

Cointegrating coefficient estimates for the basic model (4) are presented in Table 2. We introduce the estimates of (4) first, although Table 1 suggests more than one vector in some cases. The cointegrating coefficient estimates of β_1 and β_3 are in agreement with theoretical expectations. It seems that the effective federal funds rate (*FF*), and the maturity risk premium (*MRP*) as specified in (4) significantly explains variation in the U.S. Treasury yields. Expected inflation

over the next 5 to 10 years (*EI*), however, as measured by the University of Michigan's Survey of Consumers is not capturing the theorized effect on U.S Treasury yields.

Table 1. Results of Cointegration Tests.

Panel A: Basic Model						
(1) Yield	(2) Lag s	(3) Null	(4) Trace	(5) 0.05 Trace C.V	(6) Max-Eigen	(7) 0.05 Max-Eigen. C.V
i-1	4	None	51.48***	29.80	31.49***	21.13
		At most 1	19.99***	15.49	15.36**	14.26
		At most 2	4.63**	3.84	4.63**	3.84
i-2	4	None	50.41***	29.80	34.22***	21.13
		At most 1	16.19**	15.49	12.04	14.26
		At most 2	4.15**	3.84	4.15**	3.84
i-3	4	None	49.70***	29.80	33.75***	21.13
		At most 1	15.95**	15.49	12.05	14.26
		At most 2	3.90**	3.84	3.90**	3.84
i-5	4	None	54.89***	29.89	39.50***	21.13
i-7	4	None	47.89***	29.80	33.33***	21.13
i-10	4	None	65.52***	29.80	50.29***	21.13
		At most 1	16.19**	15.49		
Panel B: Composite Model with Chinese Purchases						
(1) Yield	(2) Lag s	(3) Null	(4) Trace	(5) 0.05 Trace C.V	(6) Max-Eigen	(7) 0.05 Max-Eigen. C.V
i-1	4	None	66.68***	47.86	34.65**	27.58
		At most 1	32.03**	29.80		
i-2	4	None	64.43***	47.86	38.01***	27.58
i-3	4	None	63.36***	47.86	40.87***	27.58
i-5	4	None	70.18***	47.86	51.59***	27.58
i-7	4	None	62.19***	47.86	43.29***	27.58
i-10	4	None	77.79***	47.86	51.38***	27.58

Notes: The symbols * [**] (***) attached to the figure indicate rejection of the null of no cointegration at the 10%, 5%, and 1% levels, respectively. We report the trace-statistics and maximum eigenvalue-statistics when they reject the null hypothesis listed in column (3). The lag length is chosen by the FPE, AIC, SC, or HQ Criterion. Lag-exclusion tests at the 5% helped determine the selected lag-length. The trend assumption included the linear deterministic trend.

Table 2. Cointegrating coefficient estimates

$$i_t = \beta_0 + \beta_1 FF_t + \beta_2 EI_t + \beta_3 MRP_t + \varepsilon_t. \quad (4)$$

(1)	(2) (3) (4)			(5) (6) (7)			(8) (9) (10)		
	β_1	<u>OLS estimates</u>		β_1	<u>DOLS estimates^a</u>		β_1	<u>JOH-ML estimates</u>	
		β_2	β_3		β_2	β_3		β_2	β_3
i_1	0.88*** (0.017)	-0.04 (0.077)	1.19*** (0.141)	0.88*** (0.017)	-0.05 (0.074)	0.99*** (0.122)	0.93*** (0.037)	-0.17 (0.164)	0.04 (0.263)
i_2	0.83*** (0.021)	-0.11 (0.095)	1.85*** (0.152)	0.84*** (0.023)	-0.12 (0.122)	1.77*** (0.153)	0.50*** (0.197)	-0.58 (0.888)	7.53*** (1.390)
i_3	0.78*** (0.025)	-0.130 (0.114)	2.10*** (0.169)	0.77*** (0.029)	-0.136 (0.155)	2.13*** (0.184)	0.52*** (0.146)	-0.66 (0.659)	6.40*** (1.031)
i_5	0.672*** (0.033)	-0.12 (0.142)	2.31*** (0.203)	0.67*** (0.039)	-0.10 (0.193)	2.52*** (0.234)	0.47*** (0.143)	-0.86 (0.623)	6.90*** (0.998)
i_7	0.62*** (0.04)	-0.14 (0.161)	2.39*** (0.228)	0.61*** (0.045)	-0.11 (0.215)	2.70*** (0.258)	0.39*** (0.158)	-0.92 (0.694)	7.75*** (1.107)
i_{10}	0.56*** (0.044)	-0.07 (0.172)	2.45*** (0.242)	0.54*** (0.050)	-0.01 (0.226)	2.94*** (0.301)	0.32** (0.162)	-0.69 (0.713)	8.11*** (1.137)

Notes: The dependent variables are the U.S Treasury yields. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors are reported in parenthesis for both OLS and DOLS. The symbols * [**] (***) attached to the figure indicate rejection of the null of no cointegration at the 10%, 5%, and 1% levels, respectively. ^A One lead and lag of the first-differenced FF, EI, and MRP are included in the DOLS regressions.

Table 3 presents the estimation of (5) for alternative maturities. Purchases of U.S Treasuries by China significantly lower the U.S Treasury Yield Curve. In general, an increase in the purchase of U.S. Treasuries by China leads to a significant reduction in the U.S. Treasury yields, especially the yields on the mid to long term securities such as the i_2 , i_3 , i_5 , i_7 , i_{10} Treasury yields. This seems logical: as the amount of Treasury securities purchased by China goes up, their price goes up as well and their yield comes down, *ceteris paribus*. Note that the effect of purchases of U.S Treasuries by China is stronger as the term of the security increases. For example, a one percent increase in purchases of U.S. Treasuries by China lowers the i_2 Treasury Constant Maturity Rate yield by, on average, 10 to 38 basis points (*ceteris paribus*) while a one percent increase in purchases of U.S Treasury Constant Maturity securities by China lowers the i_{10} Treasury Constant Maturity Rate yield by, on average, 39 to 55 basis points.

Table 3. Coefficient estimates of the composite model,

$$i_T = \beta_0 + \beta_1 FF_t + \beta_2 EI_t + \beta_3 MRP_t + \beta_4 \log(PUTC_t) + \varepsilon_t. \quad (5)$$

(1)	(2)	(3) (4) (5) <u>OLS estimates</u>			(5)	(6) (8) (9) <u>DOLS estimates</u> ^a				(10)	(11) (12) (13) <u>JOH-ML estimates</u>				(14)
	β_1	β_2	β_3	β_4	β_1	β_2	β_3	β_4	β_1	β_2	β_3	β_4	α		
i_1	0.88*** (0.02)	-0.04 (0.07)	1.20*** (0.14)	-0.01 (0.02)	0.88*** (0.02)	-0.05 (0.07)	0.96*** (0.14)	-0.01 (0.03)	0.85*** (0.05)	-0.07 (0.18)	0.53* (0.33)	-0.15*** (0.06)	-0.011 (0.035)		
i_2	0.67*** (0.02)	-0.08 (0.09)	1.70*** (0.15)	-0.09*** (0.03)	0.78*** (0.02)	-0.10 (0.10)	1.51*** (0.17)	-0.12*** (0.04)	0.66*** (0.07)	-0.14 (0.24)	0.51 (0.44)	-0.38*** (0.08)	-0.069*** (0.030)		
i_3	0.69*** (0.03)	-0.08 (0.114)	1.85*** (0.16)	-0.16*** (0.03)	0.68*** (0.03)	-0.10 (0.13)	1.71*** (0.19)	-0.18*** (0.04)	0.53*** (0.06)	-0.13 (0.22)	0.24 (0.41)	-0.44*** (0.8)	-0.118*** (0.033)		
i_5	0.52*** (0.03)	-0.04 (0.12)	1.88*** (0.18)	-0.28*** (0.05)	0.51*** (0.04)	-0.04 (0.15)	1.86*** (0.22)	-0.29*** (0.06)	0.34*** (0.08)	0.56* (0.29)	2.66*** (0.54)	-0.47*** (0.10)	-0.067*** (0.024)		
i_7	0.45*** (0.03)	-0.04 (0.13)	1.87*** (0.19)	-0.34*** (0.05)	0.43*** (0.04)	-0.03 (0.16)	1.92*** (0.22)	-0.35*** (0.06)	0.24*** (0.09)	0.58* (0.34)	3.45*** (0.62)	-0.51*** (0.12)	-0.049*** (0.020)		
i_{10}	0.35*** (0.04)	-0.05 (0.13)	1.83*** (0.19)	-0.39*** (0.06)	0.33*** (0.04)	0.09 (0.17)	2.03*** (0.23)	-0.40*** (0.06)	0.14* (0.09)	0.37 (0.33)	3.64*** (0.62)	-0.55*** (0.12)	-0.045*** (0.019)		

Notes: The dependent variables are the U.S Treasury yields. Newey-West heteroskedasticity and autocorrelation consistent (HAC) standard errors are reported in parenthesis for both OLS and DOLS. The symbols * [**] (***) attached to the figure indicate rejection of the null of no cointegration at the 10%, 5%, and 1% levels, respectively. ^aOne lead and lag of the first-differenced FF, EI, MRP and log(PUTC) are included in the DOLS regressions. The speed of adjustment (α) measures the impact of lagged one period deviations from the long-run vector on the yield rates differences as the dependent variable.

Based on column (14) of Table 3 for the VECM associated with the Johansen estimates, the error-correction terms (speed of adjustments) are all negative and statistically significant, except for i_j . The null hypothesis of these speeds of adjustment being zero can be rejected at conventional significance levels. When deviations from the long-run equilibrium occur it is primarily the yields that adjust to restore long-run equilibrium over our sample, rather than the included predictors. This would imply that if the U.S yield curve was higher than expected a priori in the last month, in the current month it would be decreased by, on average, 6 percent to restore the long-run relationship between the yield curve and the included determinants. In other words, the included determinants are (*weakly*) exogenous. Further, unidirectional Granger causality going from the predictors to the yield curve is supported in two ways: First, in the long-run the cointegrating coefficients are driving the yield curve with no feedback. Second, the temporal deviations from the long-run path are corrected by changes in the yield curve.

The U.S Treasury Yield Curve is not only lowered by purchases of U.S. Treasuries by China, but also flattened. Figure 5 supports this. Figure 5 was derived by taking a linear average of the extreme points of the range of results obtained from OLS, DOLS, and JOH. For example, for i_{10} the range was -39bp to -55bp for an increase of 1% in PUSTC, the mid-point of that range is 47bp, and so on. Using the observations for 2008:05, the complete series was calculated as presented. In this way the figure reflects the middle of the estimation of all the methods. However, if we choose one particular method, the result will be similar because the effect of PUSTC on the yield curve gets stronger as the term to maturity increases regardless of the method. A hypothesized 1% increase in Treasury Constant Maturity Treasury Securities purchase by China significantly lowers and flattens the U.S Treasury Yield Curve, with stronger effects on longer term securities than on shorter term securities.

Meese and Rogoff (1983) compare the forecasting performance of the basic monetary model of exchange rate determination against a naïve random walk model for U.S. dollar exchange rates for several countries. Mark (1995) evaluates performance of the basic monetary model at longer horizons relative to that of shorter horizons. We adopt a similar motivation for interest-rate based studies. Using Theil's (1966) U -statistic given by (9), a value larger than one indicates that the basic model does worse than the composite model in minimizing RMSE. The upper part of Table 4 shows the in-sample Theil's U -statistic for the basic and composite models in (4) and (5). Theil's U -statistic is, except for i_1 , greater than 1 throughout and the ratio increases linearly with the term of the securities. This lends support to the hypothesis that the composite model is superior to the basic model in predicting variation in the U.S Treasury Yield Curve. We also compute and compare the Mean Square Error (MSE) in (8) for

both models and test the null hypothesis that the MSE obtained from the basic model is equal to the MSE of the composite model against the alternative hypothesis that one MSE is smaller than the other.

Table 4. Root Mean Square Errors (RMSEs) for the Basic and Composite Models for In-sample Forecasts and for One-Step Ahead, Recursive Out-of-sample Forecast Comparisons.

In-sample forecasts:

(1) Dependent Variable	(2) RMSE _B	(3) RMSE _C	(4) U ^a	(5) MSE-t ^b
i_1	0.262	0.263	0.99	-0.005
i_2	0.353	0.332	1.06	0.749
i_3	0.429	0.377	1.14	1.786**
i_5	0.583	0.467	1.25	3.632***
i_7	0.669	0.520	1.29	4.336***
i_{10}	0.742	0.553	1.34	5.090***

Out-of-sample forecasts:

(1) Dependent Variable	(2) RMSE _B	(3) RMSE _C	(4) U ^a	(5) MSE-t ^b	(6) DM ^c
i_1	0.7313	0.7390	0.99	0.057	-3.08**
i_2	0.9756	0.9187	1.06	0.399	4.09***
i_3	1.0754	0.9701	1.11	0.690	4.02***
i_5	1.0820	0.8896	1.22	1.501*	3.97***
i_7	1.1019	0.8597	1.28	2.021**	3.96***
i_{10}	1.0220	0.7318	1.40	2.618***	3.94***

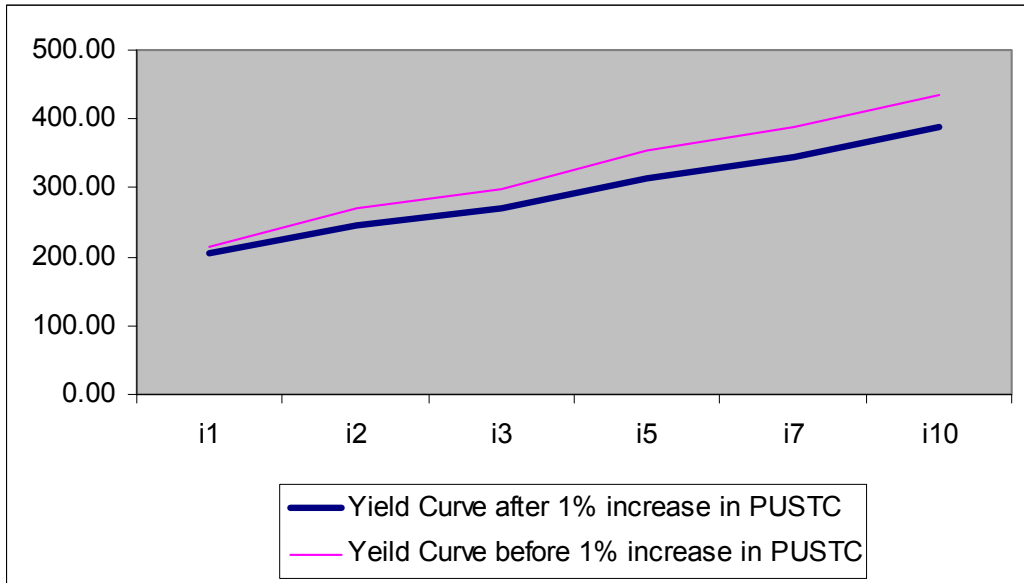
Notes: U is the ratio RMSE_B/RMSE_C, where RMSE_B is the root mean square error for the basic model and RMSE_C is the root mean square error for the composite model.

^bOne-sided (upper-tail) test of $H_0: \text{MSE}_B = \text{MSE}_C$ versus $H_1: \text{MSE}_B > \text{MSE}_C$; 10, 5, and 1 percent critical values equal 1.28, 1.64, 2.33, respectively. Negative statistics imply that the basic model forecast beats the composite model forecast. Positive statistics imply that the composite model forecast beats the basic model forecast.

^cThe Diebold-Mariano (1995) statistic (DM) is obtained by regressing the loss differential series on an intercept and a MA (1) term to correct for serial correlation. Negative statistics imply that the basic model forecast beats the composite model forecast. Positive statistics imply that the composite model forecast beats the basic model forecast.

*, **, *** indicate significant at the 10, 5, and 1 percent levels, respectively.

Figure 5. Average Effect of 1% Increase in Purchases by Chinese Investors of U.S. Treasuries on the U.S Treasury Yield Curve (left scale in Basis Points).



The lower part of Table 4 shows the one-step-ahead out-of-sample Theil's U -statistic for the basic and composite models as presented in (4) and (5). Again, Theil's U -statistic is, except for i_1 , also greater than 1 throughout and the ratio increases linearly with the term of the securities as well. As before, this lends support to the hypothesis that the composite model is superior to the basic model in predicting variation in the U.S Treasury Yield Curve even when using one-step-ahead observations to evaluate the forecasting precision of the models. We then repeat the test performed in the previous section to test that $MSE_C = MSE_B$ against the alternative hypothesis that $MSE_B > MSE_C$ when out-of-sample observations are used for prediction.

One-sided (upper-tail) t -tests of $H_0: MSE_B = MSE_C$ versus $H_1: MSE_B > MSE_C$ are presented in column 5 of the lower part of Table 4. The null hypothesis is rejected at conventional levels for the mid to long term securities. However, the null can not be rejected for the One-year, Two-year, and Three-year securities. Out-of-sample, the evidence is stronger for the augmented model improving forecasting power at longer maturities. The Diebold-Mariano (1995) statistic at the sixth column of the lower part of Table 4, however, strongly suggests that the composite model forecasts beat the ones from the basic model for all maturities, except the One-Year yield. Diebold and Lee (2006) find that their "Nelson-Siegel" factorization model of the U.S. Yield Curve is inferior to

the random walk when the horizon is only one period, with improving results for longer horizons. Overall, the forecasting exercises point to evidence that Chinese investors do exert a significant effect on the U.S. Yield Curve.

One-sided (upper-tail) t-tests of $H_0: \text{MSE}_B = \text{MSE}_C$ versus $H_1: \text{MSE}_B > \text{MSE}_C$ are presented in column 5 of the upper Table 4. The null hypothesis is rejected at conventional levels for the mid-to-long term securities. However, the null cannot be rejected for the One-year and Two-year securities. This is in line with the cointegration findings presented in Tables 4 and 5, which shows that the effect of the purchases of U.S. Treasuries by China on the U.S. Treasury Yield Curve is significantly stronger on the mid-to-long term securities.

In addition to the in-sample forecasts, we also compute one-step-ahead out-of-sample comparison as done by Rapach and Wohar (2002), as well as the Diebold and Mariano (1995) statistics. We test the null hypothesis that the Mean Square Error of the Composite Model (MSE_C) is equal to the Mean Square Error of the Basic Model (MSE_B) against the alternative hypothesis that $\text{MSE}_B > \text{MSE}_C$ using a recursive window to generate a series of out-of-sample forecasts, in our case, for the last twelve months of the full sample. The holdout sample encompasses the last twelve months of data observations.

5. Concluding Remarks

China is currently conjectured to exert considerable influence on the U.S. Treasury Yield Curve. We do confirm this broad assessment in this paper using monthly data from 1985 to 2008 on several grounds. The introduction of a demand-side variable capturing Chinese purchases of U.S. Treasuries does provide a clear improvement in the cointegrating properties of U.S. interest rates. An increase in the purchase of U.S. Treasuries by China leads to a significant reduction in the U.S. Treasury yields, especially in the mid-to-long term securities: as the amount of U.S. Treasury securities purchased by China goes up, the price of longer-term securities goes up, driving yields down, *ceteris paribus*.

Not only is the U.S. Treasury Yield Curve lowered by purchases of U.S. Treasuries by China, but also flattened: a hypothesized 1% increase in Treasury Constant Maturity Treasury Securities purchased by China significantly lowers and flattens the U.S. Treasury Yield Curve. Using the metaphor by Friedman (2005) and reviewed by Leamer (2007), the flat world is observed in financial flows as well with a stronger effect on longer term securities than on shorter term securities. The explanatory power of purchases of U.S. Treasuries by Chinese investors on the behavior of the U.S. Treasury Yield Curve is corroborated by several forecasting techniques.

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