Benchmark Status in Fixed-Income Asset Markets

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Abstract

What is a benchmark bond? We provide a formal theoretical treatment of this concept that relates endogenously determined benchmark status to price discovery and we derive its implications. We describe an econometric technique for identifying the benchmark that is congruent with our theoretical framework. We apply this to the US corporate bond market and to the natural experiment that occurred when benchmark status was contested in the European sovereign bond markets. We show that France provides the benchmark at most maturities in the Euro-denominated sovereign bond market and that IBM provides the benchmark in the US corporate bond market.

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1. Introduction

Euro-denominated sovereign bonds have highly correlated yields and are generally regarded as very close substitutes for each other. Corporate bonds mainly differ from each other in regard to their credit-risk but even credit risk is known to have common components (see Driessen, 2005). Where there is such extreme similarity across an asset class, market participants' risk exposure and hedging requirements can be met from a sub-set of the class. In this case liquidity usually concentrates on a small number of individual cases and we observe that these are often selected as reference points for the market as a whole. Sometimes this is explicit, as when yields displayed on electronic platforms are expressed as spreads over a reference security. With liquidity so concentrated most of the asset class becomes very illiquid and it becomes difficult to construct reliable indices that would normally be used for pricing. This creates and reinforces the need for an alternative reference point and opens the way for a single asset to take the role as benchmark.

It seems reasonable to expect that when a reference security emerges it would attract more liquidity and become the location of price-discovery for market-wide phenomena. Once a security can be identified as being a source of price-discovery in this way, network externalities can play a role in reinforcing its position. Thus, liquidity is as much an outcome as a cause of benchmark status. Yuan (2005) proposed the idea that benchmark status relates to price-discovery for the market as a whole. Yuan's definition highlights the benchmark security's sensitivity to systematic risk, as opposed to asset-specific risk, as central to its identification. We modify Yuan's model to derive a condition that identifies a single bond as the benchmark.

In common with Yuan, our model and empirical approach associates benchmark status with the price-discovery process rather than with either the level of the yield or the liquidity of the issue (see Hasbrouck, 1995, for a treatment in the context of equity markets). We contend that a single bond is most likely to emerge as a benchmark when there is an asset with a significant sensitivity to systematic variability and insignificant idiosyncratic variability. As a focus for market-wide price discovery the benchmark will improve as a reference point because it will lead what happens elsewhere in the market. The likelihood that a benchmark will arise also depends on how deficient the alternative (i.e., an index) is likely to be and on how expensive it is to trade the index. The prevalence of OTC trading will also tend to reduce the quality of an index because so many prices are hidden. And where a suitable, single-asset, benchmark candidate exists and becomes established, the constituents of a potential index may be degraded by insufficient liquidity to the point of reinforcing the benchmark.

The model is consistent with the view that the benchmark security provides an information externality to the market as a whole because it represents common movements of the entire market without noise. In essence a benchmark security arises out of the market's need for a focus for market-wide price-discovery and concentrates the aggregation of information reducing the costs of information acquisition in all markets where a security is traded against the benchmark.

Alternative views exist regarding the qualities associated with benchmark status. For example, in the case of euro-denominated sovereign bonds the benchmark has often been assumed to be the lowest yielding bond at each maturity and this has invariably been the German one (*e.g.*, Favero *et al.*, 2000, pages 25-26). The benchmark has also been interpreted to mean the most liquid security (see Blanco

2002) which is therefore most capable of providing a regular reference point for the market. But in the Euro-denominated sovereign bond market the Italian, not the German, are easily the largest and most liquid issues and yet they are seldom regarded as having benchmark status.

Krishnamurthy and Vissing-Jorgensen (2006) also attribute a convenience yield to US Treasuries which they associate with liquidity attributes and to the perceived low-risk status of Treasuries. Our analysis suggests that this convenience yield also derives from the fact that idiosyncratic innovations in Treasury yields are either infrequent or fully expected because of transparent monetary policy. Essentially, Treasuries have most of the qualities we associate with benchmark status. Thus, while liquidity and other attributes may play a role, our theory calls for a more formal approach to identification of benchmark status. Benchmark status cannot just be asserted on the basis of underlying liquidity, credit worthiness or the level of the yield.

Apart from the positive information externality there are other reasons why identifying a benchmark could be of benefit. In the context of the MiFID proposals of the European Commission's 'Financial Services Action Plan', sell-side agents are under pressure to demonstrate that they have delivered 'best-execution' on transactions in infrequently traded assets (Commission Directive 2006/73/EC). The existence of a reliable benchmark and the ability to derive a reference price that is relevant to each security is potentially of great benefit in this case. Additionally, in the increasingly important area of 'fair-value' accounting for financial instruments, the International Accounting Standards Board (IASB, 2005) sets out a conceptual framework for determining fair-values for initial recognition in financial statements. 'Level-2' in the 'measurement hierarchy' is described by IASB in the following terms: “If observable market prices for identical assets or liabilities are not available on the measurement date, fair-value shall be estimated using observable market prices for similar assets or liabilities adjusted as appropriate for differences, whenever that information is available.” Our methodology provides a basis for implementing such an approach.

An additional motivation for our study arises from the implementation of monetary policy in the euro area. One legacy of the introduction of the euro has been the growing recognition of the need to broaden the scope of open market operations. The European Central Bank currently concentrates on the swaps and repo market to implement its monetary policy. In its ‘General documentation on Eurosystem monetary policy instruments and procedures (2002)’, however, the ECB refers to the possible need for structural operations that may be required to influence the market’s liquidity position over long horizons. Our analysis has a bearing on the choice of policy instrument.

Finally, if benchmark status increases the liquidity of an issue and provides convenience to market participants then it will attract a convenience premium. This is attractive for the issuer seeking to reduce the cost of borrowing. Our analysis highlights those aspects of policy that can aid in the acquisition of benchmark status. The minimization of idiosyncratic innovations, rather than the augmentation of the systematic component of yields, is what matters most. In the sovereign bond case, this would require an issuer to produce as few country-specific policy innovations as possible. Building a reputation for stability (never surprising the market with idiosyncratic movements) allows market participants to look ahead with some confidence in the benchmark properties of an issue. There is also some benefit to be had from improving and maintaining the liquidity of an issue since this will improve the speed with which the security can reflect market-wide developments.

The existing empirical approaches to identifying the price discovery process are limited. Scalia and Vacca (1999), for example, use Granger-causality tests to determine whether price discovery occurs in the cash or futures market in Italian bonds. In the context of identifying benchmark status, however, we believe that Granger-Causality testing exhibits significant weaknesses, particularly for high-frequency transaction data with variable liquidity. First, it can be inconclusive because series can Granger-cause each other. Second, Granger-causality is about dynamics: it has nothing to say about long-run relationships between series[1].

We also acknowledge that if a reliable index is available, then there are various ways of identifying which individual security is the best representation of that index. For example, it would be straightforward to select the security that had highest measured correlation with the index. This strategy is redundant however when a reliable index is unavailable. The approach we invoke is suitable when liquidity concentrates on a small subset of the asset-class and renders any index constructed using the entire class as unreliable and redundant.

Our empirical method exploits the fact that yields are empirically indistinguishable from non-stationary processes. If there were a unique benchmark at every maturity, then we would expect that the yields of other bonds would be cointegrated with that benchmark. Indeed, there should be multiple cointegrating vectors centering on the benchmark bond. Our empirical approach relies on a result, based on Davidson (1998), that the structural nature of the cointegrating relationship between a benchmark bond and other bonds can be identified even in the context of quite a general theoretical framework.

We apply this methodology to two interesting cases where benchmark status is of topical interest due to the concentration of liquidity and the patent lack of reliable and unrepresentative indices. The first application concerns the identification of the benchmark in the euro-denominated sovereign bond market where benchmark status was uncertain following the introduction of the euro. The second application concerns benchmark status in the US corporate bond market where the TRACE system has increased post-trade transparency making it easier for a benchmark to be identified by market participants. In the case of corporate bonds, we show that the benchmark can be identified despite the presence of differential credit risk. However, to achieve this, our methodology must be modified to include information about credit-default. We use credit default swap data to assist in this application.

In the next section, we provide an explicit theoretical framework within which a benchmark security is defined. Section 3 presents the novel empirical methodology. The results from applying this to the euro-area bond market are presented in section 4. An application to the US corporate bond market is contained in Section 5. Section 6 contains concluding remarks and directions for future research.

2. Benchmark securities: a formal framework

Yuan (2005) formalises the concept of a benchmark security and we begin by adapting this definition to the non-stationary case.

Consider a fixed-income security as having a yield with the following factor structure[2]:

$$r_i = \beta_i \tilde{\gamma} + \tilde{\varepsilon}_i \qquad i = 1.....,n \tag{1}$$

where r_i is the nominal yield on the i[th] security, $\tilde{\gamma}$ is a systematic factor and β_i is security i's sensitivity to the systematic factor. $\tilde{\varepsilon}_i$ is the idiosyncratic shock.

Conventionally, factor pricing models place very little emphasis on issues of stationarity[3]. However, bond yields are typically found to be indistinguishable from non-stationary processes and so we modify Yuan's model by identifying the source of the non-stationarity as the systematic factor $\tilde{\gamma}$ which we assume is a general I(1) process[4] with innovations $\tilde{\mu}$. Consequently, all of the yields are themselves non-stationary.

The security specific shocks $\tilde{\varepsilon}_i \; \forall i = 1.....,n$ are assumed to be stationary ARMA processes:

$$\tilde{\varepsilon}_i = B_i(L)\tilde{\eta}_i \tag{2}$$

The parameters of the ARMA process $B_i(L)$ are also security-specific, and the $\tilde{\eta}_i$ are independently distributed with mean zero and constant variance σ_i^2. We also assume that $E(\tilde{\eta}_i\tilde{\mu}) = 0 \quad \forall i$.

At this point, we require the following Lemma:

Lemma 1. All pairs of security yields $\{r_i, \quad i = 1....,n\}$ *are cointegrated.* Proof: For any r_i and r_j equation (1) implies that

$$\frac{r_i}{\beta_i} - \frac{r_j}{\beta_j} = \left(\frac{\tilde{\varepsilon}_i}{\beta_i} - \frac{\tilde{\varepsilon}_j}{\beta_j}\right)$$

The right hand side is stationary by construction. The cointegrating vector is

$\left(\frac{1}{\beta_i}, -\frac{1}{\beta_j}\right)$ ■

The variance of the cointegrating residual is:

$$\mathrm{Var}\left(\frac{r_i}{\beta_i}-\frac{r_j}{\beta_j}\right)=\frac{\left\{B_i\left(L\right)\right\}^2\sigma_i^2}{\beta_i^2}+\frac{\left\{B_j\left(L\right)\right\}^2\sigma_j^2}{\beta_j^2} \tag{3}$$

We now introduce a formal definition of a benchmark security:

Definition 1: A benchmark security has no dependence on idiosyncratic risk

This definition is obviously idealised and immediately leads to the following

$$r_b=\beta_b\tilde{\gamma} \tag{4}$$

where the subscript *b* refers to the benchmark.

This is the idea that, where a common non-stationary factor exists and there exists a security which has sensitivity (not necessarily unitary) to this factor and little or no idiosyncratic innovations, then the conditions for the emergence of a benchmark are in place. The key element for the emergence of the benchmark is the absence (or insignificance) of idiosyncratic risk.

The next lemma follows directly from Lemma 1: .

Lemma 2. All security yields $\{r_i,\ \ i=1....,n\}$ *are pairwise cointegrated with the benchmark yield* r_b.

Proof:

From equations (1) and (4),

$$\frac{r_i}{\beta_i}-\frac{r_b}{\beta_b}=\frac{\tilde{\varepsilon}_i}{\beta_i}$$

The right hand side is stationary by assumption. The cointegrating vector is

$\left(\frac{1}{\beta_i},-\frac{1}{\beta_b}\right)$ ■

The variance of the cointegrating residual is:

$$\operatorname{Var}\left\{\frac{\varepsilon_i}{\beta_i}\right\}=\frac{\left\{B_i\left(L\right)\right\}^2\sigma_i^2}{\beta_i^2} \tag{5}$$

We can now state the main result:

Theorem 1:

The variance of the residual error in the cointegrating vector between security i's yield and any other security j=1,…,n, ≠ b, is always greater than the variance of the residual error in the cointegrating vector between security i's yield and the benchmark yield.

Proof: Compare equations (3) and (5). ■

Although the conditions set out above are quite restrictive, they are in our view essential for the emergence of a benchmark security. In particular, if a security has benchmark potential it must be the case that it is not excessively compromised by non-systematic noise. If an asset with these properties cannot be identified then it is unlikely that a benchmark will emerge. The benchmark must also become central to price-discovery or at least be quick and accurate in reflecting market-wide phenomena. These requirements suggest why benchmark status is occasionally uncertain and contested. This is due to the arrival of unexpected idiosyncratic shocks that render the existing benchmark unreliable as a pricing reference and undermine confidence in its expected future relevance. Although it is tempting to associate the benchmark thus defined with a low yielding bond this is not necessarily the case since β_b could be relatively high. In the euro-denominated sovereign bond market, for example, an absence of idiosyncratic innovations would suggest that the benchmark is likely to be associated with a very central player in the monetary union or one that is very committed to the union.

The above analysis has made no reference to credit default risk. This is hardly a significant omission in the cases of the sovereign markets of the euro-zone or the US. However, credit default is a serious issue for the corporate bond market. In particular, it would be heroic to assume that the non-stationary component of asset-specific default and credit risk could be entirely represented by sensitivity to a single non-stationary systematic factor. While the work of Driessen (2005) shows that there is evidence of systemic credit risk across corporate bonds, there is no suggestion in the literature that this represents all systematic risk or that all idiosyncratic non-stationary credit risk could be derived by way of sensitivity to this common component alone. However, idiosyncratic (perhaps non-stationary) credit risk can be observed from corporate bond 'credit default swap' (CDS) markets and we propose to use this information to control for asset-specific factors in our empirical application to US corporate bonds.

The benchmark derived, having controlled for these observable CDS spreads, can be interpreted as best representing market-wide price-discovery with regard to the remaining non-stationary common factor. This could be associated with non-default components of the corporate bond spread as described by Krishnamurthy and Vissing-Jorgensen (2006) and identified by a number of other recent contributions to the corporate bond-pricing literature including Collin-Dufresne, Goldstein, and Martin (2001), Huang and Huang (2001), and Longstaff, Mithal, and Neis (2006). In the case of Krishnamurthy and Vissing-Jorgensen, the variation in the corporate spread is the convenience yield on Treasuries but, since we consider yields rather than yield spreads, in our case it can be considered to encompass both the convenience yield of treasuries and other non-default components.

3. Econometric Methodology

The factor definition of a benchmark in Section 2, along with Lemmas 1 and 2 and Theorem 1, suggests that the benchmark should be identified from an analysis of the cointegration properties of the yield series. If a particular security is the benchmark at a given maturity, then there should be two cointegrating vectors in a three-variable system of security yields. The usefulness of this however is compromised by the presence of an identification problem. Even if we are satisfied that such cointegration vectors exist, we cannot draw any immediate conclusion about the structure of the relationships between yields, such as the identity of the benchmark, because any linear combination of multiple cointegrating vectors is itself a cointegrating vector.

A parallel development in non-stationary econometrics due to Davidson (1998) and developed by Barassi, Caporale and Hall (2000a,b) provides the empirical analogue of the theory laid out in Section 2 above. This approach tests for *irreducibility* of cointegrating relations and ranks them according to the criterion of minimum variance. The interesting feature of this method is that it allows us to learn about the structural relationship that links cointegrated series from the data alone, without imposing any arbitrary identifying conditions. In this case, the 'structural' relationship is simply the identity of the benchmark in a set of bond yields.

There is a risk of confusion in the use of the word 'structure', which has many different uses in the literature. In Davidson's approach, it refers to parameters or relations that have a direct economic interpretation and may therefore satisfy restrictions based on economic theory. It need not mean a relationship that is regime-invariant. The possibility that "incredible assumptions" (Sims, 1980) need not always be the price of obtaining structural estimates turns out to be a distinctive feature of models with stochastic trends.

According to Davidson (1998) an irreducible cointegrating vector can be defined as follows:

Definition 2 (Davidson): A set of I(1) variables is called irreducibly cointegrated (IC) if they are cointegrated, but dropping any of the variables leaves a set that is not cointegrated.

Following from this definition, IC vectors can be divided into two classes: *structural* and *solved*. A structural IC vector is one that has a direct economic interpretation. More specifically:

Theorem 2 (Davidson). If an IC relation contains a variable which appears in no other IC relation, it is structural.

The less interesting solved cointegrating vectors are defined as follows:

Definition 3 (Davidson). A solved vector is a linear combination of structural vectors from which one or more common variables are eliminated by choice of offsetting weights such that the included variables are not a superset of any of the component relations.

Thus, a solved vector is an IC vector which is a linear combination of structural IC vectors. Once an IC relation is found, we wish to distinguish between structural and solved forms. The key point is that we can identify the structure from the data directly.

The BCH extension of Davidson's framework can be illustrated concretely as follows. In a system made up of three I(1) variables, say the French, German and Italian bond yields, consider the case where the German-French yield pair and the German-Italian yield pair are both cointegrated. It follows necessarily that the French-Italian pair is also cointegrated. The cointegrating rank of these three variables is 2, and one of these three IC relations necessarily is solved from the other two. The problem is that we cannot know which. The BCH methodology gives an (almost) unambiguous answer. In order to detect which of the cointegrating relations is the solved one and which of the vectors are irreducible and structural, we calculate the descriptive statistics of each cointegrating relation and rank these vectors on the basis of the magnitude of the variance of their residual errors. The structural vectors emerge as those corresponding to the lowest variance. The reason for this is suggested by standard statistical theory and can be illustrated as follows. Let *G*, *F* and *I* be our cointegrated series. Suppose G provides the benchmark so that the first two of the following equations can be considered 'structural relations' while the third is 'solved'.

$$\begin{aligned} F &= \beta_1 G + e_1 \\ I &= \beta_2 G + e_2 \\ I &= \beta_2 F + e_3 \end{aligned} \tag{6}$$

Now e_1 and e_2, being the structural error terms are assumed[5] to be distributed independently $\mathrm{N}\left(0,\sigma_{\mathrm{i}}^2\right)$, $i = 1,2$. The third equation is just solved from the first two. This implies that e_3 is a function of e_1 and e_2, and therefore we expect it to be distributed $\mathrm{N}\left(0,\frac{\beta_2^2}{\beta_1^2}\sigma_1^2 + \sigma_2^2\right)$. Note that this is always greater than σ_2^2 and if $\beta_1 \leq \beta_2$ then $\sigma_3^2 > Max\left\{\sigma_1^2,\sigma_2^2\right\}$. The solved variance can only be smaller than the structural if $\sigma_2^2 < \sigma_1^2$ and if $\frac{\beta_2^2}{\beta_1^2}$ is small enough such that $\frac{\beta_2^2}{\beta_1^2}\sigma_1^2 < \left(\sigma_1^2 - \sigma_2^2\right)$. For this to be empirically significant, we require that either β_1 or β_2 diverge considerably from unity. Therefore cointegrating relations whose residuals display lower variance should almost always be the structural ones, the remaining others being just solved cointegrating relations. This result simply mirrors the statement of Theorem 1 in Section 2.

4. Application: Contested benchmark status in the Euro-denominated sovereign bond market

The introduction of the euro on 1 January 1999 eliminated exchange risk between the currencies of participating member states and thereby created the conditions for a substantially more integrated public debt market in the euro area. The euro-area member states agreed that from the outset, all new issuance should be in euro and outstanding stocks of debt should be re-denominated into euro. As a result, the euro-area debt market is comparable to the US treasuries market both in terms of size and issuance volume (see Galati and Tsatsaronis 2001; Blanco 2001).

Despite the homogenizing effects of the euro in terms of exchange rate risk, public debt management in the euro area remained decentralised under the responsibility of 12 separate national agencies. This decentralised management of the euro-area public debt market is one reason for cross-country yield spreads. McCauley (1999) draws some comparisons between the US municipal bond market and the euro government bond markets. But the evidence of differentiation across countries has not been thoroughly explored (see Codogno *et al*. 2003; Portes 2003). It is clear from

Blanco (2001), however, that in the initial years of the euro, yields were normally lowest for German bonds; that there was an inner periphery of countries centred on France for which yields were consistently higher; and that the outer periphery centred on Italy displayed the highest yields.

The significant levelling of the playing field brought about by the introduction of the euro was expected to impact on concentration in trading activity across the more liquid and lower default-risk bonds or in the bonds that were deliverable on futures contracts. Since much of the trading in these bonds remains OTC it is difficult to assess how concentrated liquidity was. The recent work by Dunne et al. (2006) indicates that liquidity has been surprisingly well spread within and across country-specific bonds on the MTS inter-dealer trading platform. This is largely due to the use of Primary Dealer Systems (PDS) by the issuers. A significant element of the PDS is the imposition of obligations on the primary dealers to provide liquidity on the MTS trading system.

Nevertheless, this is a relatively recent development and, as described by Galati and Tsatsaronis (2001), there was a period of competition between issuers over benchmark status at each maturity at the inception of the monetary union. While the existence of a very liquid futures market in German bunds favoured German bonds as benchmark, the fact that these bonds were not issued in great size at short and medium maturities introduced a significant element of uncertainty regarding benchmark status at these maturities. We now examine which country empirically appears to provide the benchmark in this period of contested benchmark status.

4.1 Data

We have a unique data set from Euro-MTS for 1 April 2003 to 31 March 2005. We use daily data for this period. Since the creation of the euro in 1999, Euro-MTS has emerged as the principal electronic trading platform for bonds denominated in euros. Already by the end of 2000, it was handling over 40% of total transactions volume (Galati and Tsatsaronis, 2001). Government bonds traded on Euro-MTS must have an issue size of at least €5 billion. For a discussion of MTS, see Scalia and Vacca (1999) and DeJong et al. (2004). It is possible to obtain a clearer and more recent estimate of the MTS share of the market by reference to Dunne et al. (2006). In the case of the Italian market, all trading is legally required to be conducted on the MTS trading platform so our Italian coverage is 100% of the trading in Italian bonds.

The Italian market provides a useful guide as to the likely overall trading volume in non-Italian bonds. The study by Dunne et al. (2006) shows that, in December 2003, the MTS traded volume as a percentage of the amount outstanding for the three most recently issued 10 year Italian bonds, ranged between 12.19% and 34.37%. The highest percentages for the equivalent French and German 10 year bonds were 2.74% and 3.76% respectively. Thus, even if the actual turnover in French and German bonds was as low as 12.19%, this implies MTS shares of overall trading equal to 23% in the case of France and 30% in the case of Germany. While this implies that the French and German bonds are not well represented on MTS it is the case that this platform is virtually the only, and certainly the most visible, source of pre- and post-trade information in this market and we contend that it represents the only viable location for a benchmark.

We selected data from France, Germany and Italy. Together the three countries account for over 70% of the euro-zone market (Blanco, 2001). We found that the coverage of the data for the other euro-zone countries was too sparse to get a consistently clear picture of even daily activity. In the analysis below we examine bonds from within four maturity buckets[6]. Within each maturity, we select the bond that is most recently issued for each country. This rule is applied on a 3-month basis, so that the bond used in the analysis can change a number of times over the two years of our study. At the long and very long maturities, for each country, the same bond was both available and liquid throughout the two-year sample. For the short and medium maturities, there were frequent changes of bonds, both because bonds changed maturity bracket or went off-the-run. In our empirical work we included shift dummies to account for bond changes of this kind. The daily observation for each country and maturity was chosen as the final mid-quote at the 16.30 close. We used the mid-quote to avoid spurious bid-ask bounce. It is important to emphasise that the quotes are not indicative but executable.

4.2 Results

For each maturity, each bond yield is subjected to a Dickey-Fuller test for stationarity. The results are reported in Table 1. The outcome of the tests is simple to summarise. In every case, non-stationarity[7] of the yield cannot be rejected.

In this light, our empirical strategy is as follows. First, we use the Johansen[8] procedure to identify the number of cointegrating vectors at each maturity in our three-variable system. Then, we use Phillips-Hansen fully modified estimation to estimate the irreducible cointegrating vectors as recommended by Davidson. Finally we rank the irreducible cointegrating vectors using the variance-ranking criterion of BCH. From this we identify the structural vectors and therefore the benchmark. The latter must be the common yield in the two structural irreducible cointegrating vectors. The results are shown for each maturity in Tables 2 to 5.

(i) <u>Johansen Procedure</u>**:**

In Tables 2a, 3a, 4a and 5a, it is clear that there are two cointegrating vectors among the three yields at the short, medium, long and very long maturities. This can be read from comparing the 'λ-max' and 'trace' test statistics to their critical values which we reproduce for convenience. Consequently, we conclude that all yields are pairwise cointegrated.

(ii) <u>Irreducible cointegration vectors and BCH minimum variance ranking:</u>

Regressions for each pair of yields were carried out using Phillips Hansen Fully Modified Estimation. To ensure that the choice of dependent variable does not matter we obtained estimates for each possible choice of dependent variable in each pairing. This amounts to six regressions for each maturity, 24 in all. Full details of the regressions are available on request. For each maturity, Tables 2b, 3b, 4b and 5b report the residual standard error of each regression. The results are summarised here by each maturity.

Short: The standard deviation of the residuals of the six cointegrating vectors varies from 9.10 to 30.97 basis points. The highest two arise from the regression of the Italian on the German yield and vice versa. From this we conclude that that the

French-German and Italian-French relationships are structural and that the *French yield provides the benchmark at the short end*.

Medium: The standard deviation of the residuals of the six cointegrating vectors varies from 8.78 to 12.02 basis points. As with the short maturity, the highest two arise from the regression of the Italian on the German yield and vice versa. From this we conclude that that the French-German and Italian-French relationships are structural and that the *French yield provides the benchmark at the medium maturity*.

Long: The standard deviation of the residuals of the six cointegrating vectors varies from 7.57 to 13.80 basis points. Once again, the highest two arise from the regression of the Italian on the German yield and vice versa. From this we conclude that that the French-German and Italian-French relationships are structural and that the *French yield provides the benchmark at the long end*.

Very Long: It is clear that French-German pair provides us with the lowest residual standard deviation irrespective of the choice of dependent variable. However, the range of the residual variances for the Italian-French and the German-Italian pairings overlap. This therefore is a case where the conditions for selection of the benchmark are not unambiguously met. All we can definitely conclude is that Italy does not provide the benchmark. While the assumption that France provides the benchmark is not contradicted, we cannot reach a definite ranking of pairs. From this we conclude that *benchmark status is contested at the very long maturity between France and Germany*.

5. Application: The search for a benchmark in the US Corporate bond market

In this section we consider the identification of a single-asset benchmark in the US corporate bond context. The purpose of this example is to demonstrate that the technique can be expanded to a wider range of assets so long as it is possible to account for idiosyncratic, non-stationary components. In the case of corporates, the asset-specific effects can be rolled-into the analysis by using credit default swap spreads.

The earlier methodology need only be adjusted by the inclusion of exogenous variables representing idiosyncratic non-stationary components (CDS spreads). Having controlled for these variables we expect corporate yields to be pair-wise cointegrated. We also expect the pairings with the benchmark asset to possess the lowest residual variance from the cointegrating regressions. In this case we do not consider Johansen ML regressions since this would not really provide us with useful information (i.e., we do not necessarily have priors to test regarding the cointegrating rank between all the corporate yields and their CDS spreads). What we propose instead is to consider whether corporate yields are non-stationary and whether they are pair-wise cointegrated conditional on CDS spreads. Once this is established we can proceed to assess which of the assets appears to provide the benchmark based on the minimum variance technique.

5.1 Corporate bond and CDS Data

There are many agencies that identify a sub-class of corporate bonds as, in some sense, providing a benchmark (e.g., JPMorgan, US Liquid Corporate Index and CreditTrade's Benchmark American Corporates). These generally make some reference to the liquidity of the constituents and/or their investment grade. We take CreditTrade's list of US benchmarks as a starting point in the search for a 'price-discovery-based' benchmark. We focus on a narrow industry grouping (manufacturing/technology) and on a maturity category (roughly 10 years to maturity). In this industry category there were 14 corporate entities classified as benchmarks. Of these only 4 were in the 10 year to maturity category and we dropped one of these because it was not frequently traded (as revealed by the TRACE system). This provides us with a corporate analogue to the sovereign case analyzed above (details regarding these corporates is given in the appendix).

For the 3 corporate entities chosen as potential benchmarks we obtained daily data from the Thomson Financial DATASTREAM research database for a period running from 21 March 2005 to 17 Oct 2006. The corporate yield data is sourced from FT Interactive Data and this provides an end-of-day 'evaluated price' in each case. The CDS data is sourced from CMA records and provides an implied end-of-day yield spread based on the average mid-price between bid-offer quotes for CDS premia collected from a number of CDS dealers[9].

5.2 Results

ADF stationarity test results for the corporate yields are shown in Table 6a. This clearly indicates that non-stationarity cannot be rejected for all of the yields[10]. What is more interesting is the evidence of non-stationarity of the yields when adjusted by subtraction of the implied credit risk premium as shown in Table 6b. Credit risk is not therefore the only source of non-stationarity in corporate yields and therefore it is of interest to consider whether what is left-over contains a common non-stationary component. Although we could proceed with yields adjusted for implied credit risk we chose to include all yields and CDS spreads individually in the cointegration analysis.

Our pair-wise cointegrating relations were therefore of the form;

$$Yield_A = \beta\, Yield_B + \lambda\, CDS_A + \delta\, CDS_B + \phi\, CDS_C \tag{7}$$

In all cases the regression was conducted using the Phillips-Hansen Fully Modified estimation technique. The regression residuals were tested for stationarity and the results of these tests are presented in Table 7a. In all cases we can conclude that there is evidence of pair-wise cointegration.

To ensure that the choice of dependent variable did not change our conclusions we conducted the Phillips-Hansen regressions with every possible configuration of the pairings (i.e., 6 regressions in all). We then calculated the standard deviation of the residuals from these regressions and applied the minimum variance criterion. The relevant standard deviations are provided in Table 7b. Irrespective of how the pairings are regressed the conclusion remains the same. On the basis of this we can conclude that IBM is potentially the benchmark in this market. In view of our theory this implies that IBM corporate bonds must be liquid enough to provide timely reaction to market-wide events. Therefore IBM is a natural reference point because of the diversity of its activities or have become one due to acquisition of a market-wide price discovery role. IBM also must have few innovations that are purely idiosyncratic.

6. Conclusion

We focus on the meaning of 'benchmark' bond and consider its role in pricing when there is concentration of liquidity and price discovery. We show that the Modified Davidson Method is an econometric technique that enables us to identify the benchmarks in the European bond market and that a slight enhancement of the technique enables it to be extended to the case of the US corporate bond market.

The simple idea that the security with the lowest yield provides the benchmark has no direct role to play in the analysis although we do acknowledge that having lower idiosyncratic risk, being liquid and being a source of price-discovery for the market could attract a convenience premium and therefore give rise to a lower yield. Instead it is the lack of idiosyncratic risk that initially provides the conditions for a single security to emerge as a benchmark. Once established, the benchmark will acquire additional liquidity and have high information content for the market as a whole.

Our analysis undermines the conventional view of Germany as the benchmark issuer in the euro-denominated sovereign bond markets. What is striking is that France appears to dominate at all but the longest maturity[11]. This is consistent with the view that France has been the source of very few country-specific innovations to yields. Our identification of IBM as a corporate benchmark is perhaps not surprising given its enormous footprint in the general area of manufacturing/technology and the diversity of its business activities, but this amounts to concluding that exposure to systematic risk is what determines benchmark status. We contend that the yield on IBM corporate bonds must also have possessed a very low variance in its idiosyncratic component.

Both the theoretical framework and the econometric methodology presented here are completely general and are not specific to the particular applications offered as illustrations. Whenever a market displays a concentration of liquidity (and a degree of opaqueness in prices of potential index constituents), a benchmark asset is likely to arise as a way of providing price-discovery convenience. This is true in most bond markets. More generally, we believe that the analysis is applicable to any market where information acquisition is significantly concentrated.

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Table 1: ADF Stationarity test results.

SERIES	t-value	Conclusion
SHORT YIELD		
Italian	-0.787	Non-stationary
French	-1.148	Non-stationary
German	-0.776	Non-stationary
MEDIUM YIELD		
Italian	-0.861	Non-stationary
French	-0.813	Non-stationary
German	-0.880	Non-stationary
LONG YIELD		
Italian	-0.959	Non-stationary
French	-0.953	Non-stationary
German	-0.920	Non-stationary
V-LONG YIELD		
Italian	-1.001	Non-stationary
French	-1.049	Non-stationary
German	-1.019	Non-stationary

Table 1 contains results of Dickey-Fuller style stationarity tests on representative Italian, French and German bond yields at four different maturities. The Dickey Fuller Test was used in each case with neither trend nor lags of differences. In the cases of the short and medium yields, dummies were included to account for changes in bonds as they matured.

Table 2a: Cointegrating rank: Short Maturity

Johansen Test of Cointegrating Rank.					
Endogenous Variables: Italian, French and German Yields.					
Exogenous variables in cointegration space: Dummies for changes in bonds					
No constant inside or outside cointegration space.					
Lag length: 2	Effective sample size: 498				
Eigenvalues	λ-max	Trace	H0: rank	λ-max Crit. 90%	Trace Crit. 90%
0.7220	638.81	700.2	0	11.23	21.58
0.1154	61.20	61.35	1	7.37	10.35
0.0003	0.16	0.16	2	2.98	2.98
Conclusion					
Both the λ-max and Trace statistics imply that there are two cointegrating vectors.					

Table 2b: Irreducible Cointegration Short Maturity

Philips-Hansen pair-wise cointegrating regressions:					
Standard Deviation of residuals (in basis points) for all regression pairings. Dependent variable listed first – all regressions include intercept, trend and bond-change dummies. Lag window used in Fully-Modified estimation = 2.					
French-German	German-French	Italian-French	French-Italian	Italian-German	German-Italian
25.08	21.98	9.10	9.29	30.97	27.21
Conclusion:					
French bond provides the benchmark.					

Table 2a contains results of Johansen tests to determine the cointegrating rank in the three variable system containing Italian, French and German bond yields at the short maturity. For the same maturity, Table 2b presents the statistics for the standard deviation of the residuals from the cointegrating regressions associated with all the possible pairings of the Italian, French and German bond yield series.

Table 3a: Cointegrating rank: Medium Maturity

Johansen Test of Cointegrating Rank.					
Endogenous Variables: Italian, French and German Yields.					
Exogenous variables in cointegration space: Drift and dummies for changes in bonds					
Unrestricted constant outside cointegration space.					
Lag length: 2	Effective sample size: 498				
Eigenvalues	Λ-max	Trace	H0: rank	λ-max Crit. 90%	Trace Crit. 90%
0.8566	969.08	1319.5	0	16.13	39.08
0.4954	341.33	350.38	1	12.39	22.95
0.0180	9.05	9.05	2	10.56	10.56
Conclusion					
Both the λ-max and Trace statistics imply that there are two cointegrating vectors.					

Table 3b: Irreducible Cointegration: Medium Maturity

Philips-Hansen pair-wise cointegrating regressions:					
Standard Deviation of residuals (in basis points) for all regression pairings. Dependent variable listed first – all regressions include intercept, trend and bond-change dummies. Lag window used in Fully-Modified estimation = 2.					
French-German	German-French	Italian-French	French-Italian	Italian-German	German-Italian
8.81	8.78	10.16	10.09	12.02	11.81
Conclusion:					
French bond provides the benchmark.					

Table 3a contains results of Johansen tests to determine the cointegrating rank in the three variable system containing Italian, French and German bond yields at the medium maturity. For the same maturity Table 3b presents the statistics for the standard deviation of the residuals from the cointegrating regressions associated with all the possible pairings of the Italian, French and German bond yield series.

Table 4a: Cointegrating rank: Long Maturity

Johansen Test of Cointegrating Rank.					
Endogenous Variables: Italian, French and German Yields.					
Exogenous variables in cointegration space: Drift.					
Unrestricted constant outside cointegration space.					
Lag length: 2	Effective sample size: 498				
Eigenvalues	λ-max	Trace	H0: rank	λ-max Crit. 90%	Trace Crit. 90%
0.0602	30.98	60.44	0	16.13	39.08
0.0511	26.19	29.46	1	12.39	22.95
0.0065	3.27	3.27	2	10.56	10.56
Conclusion					
Both the λ-max and Trace statistics imply that there are two cointegrating vectors.					

Table 4b: Irreducible Cointegration: Long Maturity

Philips-Hansen pair-wise cointegrating regressions:					
Standard Deviation of residuals (in basis points) for all regression pairings. Dependent variable listed first – all regressions include intercept and trend. There are no bond-changes. Lag window used in Fully-Modified estimation = 2.					
French-German	German-French	Italian-French	French-Italian	Italian-German	German-Italian
7.62	7.57	11.83	11.33	13.80	13.13
Conclusion:					
French bond provides the benchmark.					

Table 4a contains results of Johansen tests to determine the cointegrating rank in the three variable system containing Italian, French and German bond yields at the long maturity. For the same maturity Table 4b presents the statistics for the standard deviation of the residuals from the cointegrating regressions associated with all the possible pairings of the Italian, French and German bond yield series.

Table 5a: Cointegrating rank: Very-Long Maturity

Johansen Test of Cointegrating Rank.					
Endogenous Variables: Italian, French and German Yields.					
Exogenous variables in cointegration space: Drift.					
Unrestricted constant outside cointegration space.					
Lag length: 2	Effective sample size: 498				
Eigenvalues	λ-max	Trace	H0: rank	λ-max Crit. 90%	Trace Crit. 90%
0.1581	85.72	125.0	0	16.13	39.08
0.0702	36.23	39.27	1	12.39	22.95
0.0061	3.04	3.04	2	10.56	10.56
Conclusion					
Both the λ-max and Trace statistics imply that there are two cointegrating vectors.					

Table 5b: Irreducible Cointegration: Very-Long Maturity

Philips-Hansen pair-wise cointegrating regressions:					
Standard Deviation of residuals (in basis points) for all regression pairings. Dependent variable listed first – all regressions include intercept and trend. There are no bond-changes. Lag window used in Fully-Modified estimation = 2.					
French-German	German-French	Italian-French	French-Italian	Italian-German	German-Italian
7.71	7.62	12.26	11.47	13.16	12.18
Conclusion:					
Benchmark status is contested between France and Germany					

Table 5a contains results of Johansen tests to determine the cointegrating rank in the three variable system containing Italian, French and German bond yields at the very-long maturity. For the same maturity Table 5b presents the statistics for the standard deviation of the residuals from the cointegrating regressions associated with all the possible pairings of the Italian, French and German bond yield series.

Table 6a: ADF Stationarity Tests
Unadjusted Corporate Yields.

SERIES	t-value	Crit. 95%	Conclusion
10 year YIELD			
Maytag	-2.276	-2.87	Non-stationary
Paper	-1.266	-2.87	Non-stationary
IBM	-1.083	-2.87	Non-stationary

Table 6a contains results of Dickey-Fuller style stationarity tests on corporate bond yields of Maytag Corporation, International Paper and IBM. The (Augmented) Dickey Fuller Test was used in each case without trend but with lags of differences equal to 2, 5 and 0 for Maytag, Paper and IBM respectively. Lags were included until Ljung-Box(Q) statistics were sufficient to reject serial correlation up to the 10th lag with 90% confidence. The number of observations before lags and differences was 412. The ADF critical value for the 95% confidence interval is given for this sample size as it is reported in the Microfit 4.1. econometric software package.

Table 6b: ADF Stationarity Tests
Corporate Yields less the Implied Credit Spread.

SERIES	t-value	Crit. 95%	Conclusion
10 yr YIELD less credit spread			
Maytag	-1.222	-2.87	Non-stationary
Paper	-0.825	-2.87	Non-stationary
IBM	-0.994	-2.87	Non-stationary

Table 6b contains results of Dickey-Fuller style stationarity tests on corporate bond yields adjusted for credit default, of Maytag Corporation, International Paper and IBM. The (Augmented) Dickey Fuller Test was used in each case without trend but with lags of differences equal to 2, 8 and 1 for Maytag, Paper and IBM respectively. Lags were included until Ljung-Box(Q) statistics were sufficient to reject serial correlation up to the 10^{th} lag with 90% confidence. The number of observations before lags and differences was 412. The ADF critical value for the 95% confidence interval is given for this sample size as it is reported in the Microfit 4.1. econometric software package.

Table 7a: ADF Stationarity Tests
Residuals from Philips-Hansen pair-wise cointegrating regressions.

SERIES	t-value	Crit. 90%	Conclusion
Regression residual			
Paper on Maytag (6 ADF lags)	-4.566	-3.07	Cointegrated
IBM on Maytag (6 ADF lags)	-5.492	-3.07	Cointegrated
IBM on Paper (7 ADF lags)	-3.093	-3.07	Cointegrated

Table 7a contains results of Johansen tests to determine the cointegrating rank in the three variable system containing the corporate yields of Maytag Corporation, International Paper and IBM. The (Augmented) Dickey Fuller Test was used in each case without constant or trend but with lags of differences equal to 2, 8 and 1 for Maytag, Paper and IBM respectively. Lags were included until Ljung-Box(Q) statistics were sufficient to reject serial correlation up to the 10^{th} lag with 90% confidence. The number of observations before lags and differences was 412. The ADF critical value for the 90% confidence interval is given for a sample size of 500 as it is reported in Table B.9 in Hamilton's (1994) 'Time Series Analysis' where we have 2 additional non-stationary right-hand variables in the cointegrating relation[12].

Table 7b: Corporate Bond Yields at Long Maturity, Irreducible Cointegration

Philips-Hansen pair-wise cointegrating regressions:					
Standard Deviation of residuals for all regression pairings.					
Dependent variable listed first – all regressions include intercept and CDS of all three corporates. Lag window used in Fully-Modified estimation = 5.					
IBM-Paper	Paper-IBM	IBM-Maytag	Maytag-IBM	Maytag-Paper	Paper-Maytag
0.093	0.113	0.140	0.162	0.166	0.175
Conclusion:					
IBM bond provides the benchmark.					

Table 7b presents the statistics for the standard deviation of the residuals from the cointegrating regressions associated with all the possible pairings of the corporate bond yields of Maytag Corporation, International Paper and IBM.

Appendix A.

International Paper has significant global businesses in paper, packaging and forest products. The company has operations in nearly 40 countries, employs approximately 68,700 people worldwide and exports its products to more than 120 nations. Sales of almost $24 billion annually are derived from businesses located primarily in the United States, Europe, Latin America, Asia/Pacific and Canada.

Maytag Corporation was a home and commercial appliance company, headquartered in Newton, Iowa. Sales were $4.7 billion in 2005 and the company had approximately 18,000 employees worldwide. Maytag was acquired by Whirlpool Corporation in March, 2006. Whirlpool had sales of more than $19 billion in 2005 and more than 80,000 employees worldwide.

IBM supplies IT hardware, software and Services and is also active in the delivery of financial services. In 2005 it had revenues of $91.1 billion and 329,373 employees.

1 We did in fact carry out Granger-causality tests. The results are generally inconclusive.

2 In what follows, all variables are implicitly indexed by time. To avoid cluttering the notation, we suppress the time subscripts.

3 See for example John Cochrane's book, 'Asset Pricing' (2000). It devotes less than two pages to stationarity.

4 This structure can be motivated in the sovereign bond case as inverse money demand functions with nonstationary velocity. For example: $r_i = -\beta_i Log\ v + \text{constant} + \text{noise}$ where v is the velocity of money. The latter is typically non-stationary. β_i is the inverse of the interest semi-elasticity of the demand for money and is country-specific.

5 If residuals from the structural vectors are not orthogonal, then it is not clear what 'structural' means in this context. It is essential one way or the other to make some assumption about the covariance between the structural relations. Any assumption other than a zero value, however, makes the application of the irreducible cointegrating vector approach inconclusive.

6 Short-dated bonds have maturities between 1.25 and 3.5 years. Medium, long and very long bonds have maturity spans of 3.5-6.5 years, 6.6-13.5 years and >13.5 years respectively. There is also a fifth category for bills: securities with maturity less than 1.25 years. However, until recently, only Italy was significantly trading such instruments on Euro-MTS.

7 It is important to emphasise that we are not claiming that *economically* yields have unit roots: intuitively one would expect yields to be mean reverting. Instead, we find statistically that yields are so long-memory that we cannot reject the hypothesis of non-stationarity.

8 Bowe and Mylonidis (1999) have previously applied the Johansen Procedure to test for integration in European bond markets. Their work relates to the pre-euro period.

9 CMA stands for Credit Market Analysis Ltd, a London based provider of financial data, see
http://www.creditma.com/products.aspx

10 See Endnote 7 above

11 For early support of this view, at least at the shorter maturities, see Jessen and Matzen (1999). See also Favero, Pagano and von Thadden (2005).

12 We found that the IBM CDS implied yield premium was stationary.