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EXCHANGE RATES, COUNTRY PREFERENCES, AND GOLD

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ABSTRACT

This paper provides indirect tests of the hypothesis that exchange rate movements may be largely coterminus with changes in preferences for holding claims on different countries. It is argued that changes in country preferences will be reflected systematically in the price of gold and, hence, that gold price movements, under the maintained hypothesis, should have explanatory power with respect to exchange rate movements over and above the effects of monetary shocks. The paper applies multivariate vector autoregression and cointegration modeling techniques to test for the short- and long-run influence of gold prices on exchange rates conditional on other monetary and real macroeconomic variables, and applies the resulting error correction exchange rate equation to out-of-sample forecasting exercises.

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I. Conceptual Background

Most exchange rate models that have emerged since the early 1970s have involved "asset market views," recognizing that international capital movements are not exogenous.[1] A widely employed taxonomy divides the asset market models into two broad classes: the monetary approach and the portfolio balance approach.[2] The key distinction is whether portfolio holders are assumed to regard domestic and foreign assets as perfect substitutes--in which case the model represents the monetary approach--or, equivalently, whether the uncovered interest parity condition (UIP) is assumed to hold continuously.[3]

For those subclasses of the monetary approach in which UIP enters the model structure explicitly, and for virtually all portfolio balance models, closed-form solutions can be derived for the expected rate of change of the exchange rate, but the current level of the exchange rate cannot be explained independently of its expected future level at some horizon. One approach to modeling expectations about future nominal exchange rates is to adopt a conventional model of expected inflation factors, and to assume that expectations about future real exchange rates are based on perceptions of some kind of goods market or balance of payments constraint.[4] The balance of payments constraint need not be precisely defined.[5] The key point, as Dooley (1982) has emphasized, is that the nature of any balance of payments constraint essentially reflects perceptions of a country's creditworthiness, and that changes in perceived creditworthiness will have predictable effects on both the tightness of the constraint and the real exchange rate consistent with meeting the constraint. The bottom line is that the expected future level or time path of a country's real exchange rate will presumably change systematically in response to revisions in expectations about the prospective relative returns on assets located in that country, or on financial claims held against that country.

Strong but casual empirical support for this line of argument is provided by observing that nominal exchange rates have shifted dramatically during episodes involving major changes in the relative attractiveness of holding assets in different countries. For example, the outbreak of the international debt crisis in 1982 led to very substantial real depreciations of the currencies of debt-burdened developing countries.[6] More rigorous empirical support has been difficult to muster, however, mainly because it is difficult to find time series of suitable quality and length on variables that indicate the relative attractiveness of holding claims on different countries.[7]

This paper attempts to fill this void by providing indirect empirical evidence based on the relationship between exchange rates and the price of gold. The basic line of argument is that changes in relative country preferences should be systematically reflected in the price of gold, which we view as "an asset without a country." Hence, if the effects of monetary shocks can be isolated, evidence that residual changes in the price of gold are capable of "explaining" residual changes in exchange rates might be regarded as indirect evidence that exchange rate behavior reflects changes in country preferences.

II. Theoretical Priors and Summary of Empirical Findings

Consider a world consisting of two countries, A and B, and three types of assets: net claims on A, net claims on B, and gold. Claims on A may be physical assets located in country A or net financial claims against the government or private sector of country A. The key distinguishing feature is that the prospective returns on these net claims depend on economic and political developments in country A. Adverse macroeconomic shocks to country A, other things equal, will reduce prospective yields on net claims on A. Similarly, political developments that portend higher tax rates on claims on A, other things equal, will reduce the relative attractiveness of these net claims.

Gold is viewed as an asset without a country. Gold can be held outside the jurisdiction of all tax authorities, and the return on gold is not subject to the country-specific uncertainties that surround claims on future output. Any type of shock that reduces the attractiveness of holding claims on A, other things equal, will increase the demands for other assets--both claims on

B and gold--leading to changes in market-clearing prices. Such adjustments will result in higher currency A prices for both gold and currency B (i.e., a depreciation of currency A against currency B). The currency B price of gold may rise or fall, depending on the relative strengths of the different substitution effects.

One way of formalizing the argument is as follows. The level of the exchange rate, s, can be thought of as being determined by two components--monetary fundamentals and country preferences:

$$s_t = \beta' x_t + z_t \qquad (1)$$

In equation (1), x_t is a vector of monetary fundamental variables, and z_t is an unobservable variable measuring the attractiveness of the home country's assets relative to those of the foreign country. The term $\beta' x_t$ may be thought of as a standard exchange rate model. Since z_t is unobservable, we suggest proxying it (or rather, proxying its inverse) by the domestic-currency price of gold. Thus, the empirically implementable counterpart to equation (1) that we suggest is:

$$s_t = \beta' x_t + \kappa g_t \qquad (2)$$

If s_t is the foreign price of domestic currency and g_t is the domestic price of gold, we expect κ to be negative--a rise in g_t, other things equal, represents a relative decline in the attractiveness of domestic assets and will therefore tend to be associated with a depreciation of the domestic currency.

The specification of equations (1) and (2) reflects the assumption that there are only two types of shocks: monetary shocks, and shocks that affect country preferences. We refer to the latter as real shocks. A deeper analysis would distinguish between different types of monetary and real shocks--in particular, between global real shocks that might increase the attractiveness of gold vis-à-vis both claims on A and claims on B, and country-specific real shocks that affect the attractiveness of gold vis-à-vis claims on A differently than the attractiveness of gold vis-à-vis claims on B. By implicitly assuming that all real shocks are country specific, our methodology appears to make it more difficult to find statistically significant evidence that movements in the price of gold contain information capable of explaining movements in the exchange rate between currencies A and B.

We think of monetary shocks as shocks that have no effects on the relative attractiveness of holding assets in different countries. These include both global and country-specific inflationary shocks accompanied by monetary policy responses that essentially hold constant the expected real yields on claims on A and claims on B. Such shocks generally lead to changes in nominal interest rates and, hence, the nominal carrying cost of gold, which in turn lead to jumps in the nominal price of gold. Since our objective is to extract from the price of gold only information that can be taken to reflect changes in country preferences, our econometric methodology must be capable of isolating movements in the price of gold that cannot be attributed to monetary shocks.

In the empirical model suggested in equation (2), the monetary fundamentals represented by x_t typically include relative money supplies, relative interest rates, and so on. Accordingly, the estimate of κ should only reflect variation in the gold price that is not common to these monetary fundamentals. This point becomes even more obvious if we apply the so-called Frisch-Waugh theorem (see e.g., Maddala, (1977), p. 462). Suppose we regress g_t onto x_t and retrieve the residuals, $\tilde{g}_t$ say:

$$\tilde{g}_t = g_t - \hat{\alpha}' x_t \tag{3}$$

where $\hat{\alpha}$ represents the least squares estimate. If we then regress s_t onto $\tilde{g}_t$:

$$s_t = \kappa \tilde{g}_t \tag{4}$$

the least squares estimate of κ obtained from estimation of (4) is identical to that obtained from least squares applied to (2).[8] Thus, by including monetary fundamentals in our estimating equations, we can isolate co-variation in exchange rates and gold that is not attributable to movements in the monetary fundamentals.

We have conducted a variety of empirical tests, focusing on end-of-month exchange rates between the U.S. dollar and four other major currencies (the pound sterling, the Japanese yen, the deutsche mark, and the French franc), as well as on the mark/yen exchange rate, over the period 1976-90. In light of the prominent attention earned by the work of Meese and Rogoff (1983a, 1983b, 1988), we begin by investigating how their general exchange rate equation specification[9] performs when the price of gold is added to the set of explanatory variables. We find that the price of gold is the most significant explanatory variable in-sample for an equation explaining logarithmic changes in the exchange rate. As a test of whether the information contained in the price of gold indeed reflects the prominent role that gold has played as an asset, we also present a set of results in which the price of gold is replaced by the price of wheat.

We then move from single-equation tests to tests based on a vector autoregression (VAR) system. This is followed by a third line of investigation, which uses the VAR system to examine the long-run relationships between the exchange rate, the price of gold, and the other variables in the system. We find that the long-run relationship between the exchange rate and the price of gold is highly significant with the anticipated sign. The estimated cointegration relationships are then used to find dynamic error correction equations, and we again apply recursive tests of predictive ability.

III. Empirical Results

1. Data

The data are taken from the International Monetary Fund's IFS data tape and run from January 1976 through December 1990. In particular, the exchange rate used is line ag; money is M1, line 34; the short-term interest rate is line 60c; consumer prices are from line 64; and the income measure is industrial production, line 66c. The money supply, price, and industrial production series are seasonally adjusted. The countries considered are the United States, the United Kingdom, France, Germany and Japan. We consider US dollar bilateral exchange rates, as well as the yen/mark rate.

2. Improving upon a standard empirical specification

We started by estimating an equation of the form:

$$\Delta s_t = \alpha_1\Delta(m-m^*)_t + \alpha_2\Delta(y-y^*)_t + \alpha_3\Delta(r-r^*)_t + \alpha_4\Delta(\pi-\pi^*)_t \tag{5}$$

where s_t is the foreign currency price of the U.S. dollar (or the mark price of yen); m, y, r and π denote, respectively, US narrow money, industrial production, short-term interest rate and inflation rate; and an asterisk denotes the corresponding foreign variable.[10]

Equation (5) is a general monetary model formulation, which nests the real interest differential formulation of Frankel (1979) $(\alpha_1<0,\alpha_2>0,\alpha_3>0,\alpha_4<0)$; a simple "overshooting" empirical model $(\alpha_1<0,\alpha_2>0,\alpha_3>0,\alpha_4=0)$; and the basic, flexible-price monetary model $(\alpha_1<0,\alpha_2>0,\alpha_3<0,\alpha_4=0)$.[11] In order to provide an indirect test of the hypothesis outlined above, we also estimated equations including (the logarithms of) the dollar price of gold and the dollar price of wheat.[12]

The results are reported in Table 1. As was expected, equation (5) performs badly as a model of exchange rate determination, with most coefficient estimates insignificantly different from zero and low overall explanatory power. Adding the gold variable to this equation, however, produces spectacular results: the gold price is, in every case, strongly significant and of the expected sign (negative), the R^2 rises by an average factor of six, and the Durbin-Watson statistic improves dramatically. Moreover, this success is not repeated when the price of wheat variable is included: it enters in each case with an estimated coefficient which is statistically insignificant from zero.

This set of tests, therefore, provides strong initial support for our conjecture that gold price movements should have explanatory power with respect to exchange rate movements, over and above the effects of monetary shocks.

3. Forecasting the exchange rate

We next tested to see whether knowledge of the gold price would enhance our ability to forecast the exchange rate. Rather than conduct this experiment within the very narrow empirical framework of the previous section, however, we chose to use a less restrictive, vector autoregressive (VAR) framework. Accordingly, we estimated sixth-order, unrestricted VARs, which included domestic and foreign money supplies, output, interest rates, and inflation, as well as the exchange rate, and used this to forecast the exchange rate dynamically from one to twelve months ahead. Starting with an initial VAR, estimated using data for 1976(1)-1988(12), we sequentially added one extra data point and forecast the exchange rate dynamically, using a Kalman filter algorithm to update the coefficient estimates as each new data point was added. Finally, we computed the root mean square error (RMSE) of the forecast at each horizon. As a benchmark for this exercise, we considered the resulting RMSEs relative to that produced by a random walk forecast; the resulting measure is known as Theil's U-statistic. If the U-statistic is less than unity, a superior performance to the random walk is indicated;

a value greater than unity implies an inferior performance to the random walk forecast.

The results are reported in Tables 2a-2e. Each of these tables reports four variants of this exercise plus two additional cases for the system excluding gold. The different variants distinguish between cases in which the VAR was used to forecast the values of the "exogenous" variables (money, inflation, output, and interest rates) and gold, and cases in which the actual values of these exogenous variables were used to forecast future values. We also distinguish between cases in which forecast rather than actual prices of gold were used in the forecasts.

For horizons of one to three months, the results are mixed, although it is possible that many of the U-statistics for these horizons are insignificantly different from unity.[13] For forecast horizons beyond three months, however, a VAR system including gold always gives the lowest U-statistic -- for all exchange rates.

Interestingly, however, better results were generally obtained when forecast rather than actual gold prices were used in the forecasts. For exchange rates not involving the mark, better results were also obtained using forecast values of the exogenous variables. For the mark/dollar and yen/dollar exchange rates, best results were obtained using the actual future values of exogenous variables, combined with the forecast values of the gold price.

We next repeated this experiment, replacing the price of gold with the price of wheat. For exchange rates not involving the mark, the lowest RMSEs were obtained using forecast values of the exogenous variables and of the wheat price; for mark exchange rates, best results were obtained when the actual future values of the exogenous variables were used. We then took the smallest RMSE for each exchange rate, and divided it by the corresponding RMSE obtained using the gold price in the VAR. The results are given in Table 3. An entry greater than unity in Table 3 indicates a superior performance of the system involving gold. For horizons of six months or more ahead, the system involving gold uniformly does better than the system involving wheat, although the superior performance of the gold price is not as striking as it is in the results reported in Table 1. Nevertheless, these results confirm the findings of the previous section, that the price of gold has explanatory power for the exchange rate, even when the monetary fundamentals are controlled for. Moreover, these results appear to be peculiar to gold rather than to commodities in general.

4. Cointegration and error correction

The final step in our analysis was to investigate the long-run relationship between exchange rates and gold--controlling for monetary variables--using cointegration techniques. We also used our estimated cointegrating vectors to form short-run, dynamic error-correction equations for the exchange rate, which were then subjected to post-sample forecasting tests. In our cointegration tests, we employed the maximum likelihood cointegration technique of Johansen (1988). To do this, we estimated a vector autoregression (VAR) for the variables in question[14] (in each case a sixth-order VAR was adequate in terms of residual diagnostics).

Let $X_t = (s_t, g_t, m_t, m_t^*, y_t, y_t^*, r_t, r_t^*, p_t, p_t^*)$. Then, excluding the constant terms for expositional purposes, the estimated VAR is of the form:

$$X_t=\sum_{i=1}^{6} \pi_i X_{t-i} + \epsilon_t \tag{6}$$

where the π_i are conformable coefficient matrices and ϵ_t is a white noise innovation vector. The long-run static equilibrium corresponding to (6) is [15]:

$$(I-\sum_{i=1}^{6} \pi_i)X_t=0 \tag{7}$$

The coefficient matrix in (7) can be factorized:

$$(I-\sum_{i=1}^{6} \pi_i)=\alpha'\beta \tag{8}$$

where α and β are 6xr matrices, r $\leq$ 9, such that each of the linear combinations of X_t formed by the rows of β is stationary, or I(0):

$$\beta_i' X_t \sim I(0) \tag{9}$$

for $\beta' = (\beta_1 \ldots \beta_r)$. Thus, if a stable long-run relationship between the elements of X_t exists, at least one of the β_i vectors must be significantly different from the null vector. If this is the case, then the elements of X_t are said to be cointegrated (Engle and Granger, 1987). Johansen (1988) develops a maximum likelihood technique for estimating α and β and testing for the number of distinct cointegrating vectors, r, as well as for testing linear restrictions on the parameters of the cointegrating vectors, β_i.

The results of applying the Johansen procedure to the data are reported in Table 4. Among the ten variables considered, there are at least six significant cointegrating vectors (seven in the case of sterling/dollar, mark/dollar, and yen/dollar). Moreover, the likelihood ratio statistic for the null hypothesis that gold should be excluded from the cointegrating vector is in every case greatly significant. This therefore indicates that gold is an important long-run determinant of the exchange rate, even controlling for the effects of real and monetary variables.

We then took the most significant cointegrating vector for each of the exchange rate groups and used this to find an error-correction form equation. According to the Granger Representation Theorem (Engle and Granger (1987)), if a set of first-difference stationary series X_t are cointegrated such that

$$e_t = \beta' X_t \sim I(0) \tag{10}$$

then there exists an error correction representation of the form:

$$\Delta X_t \Psi(L) = -\rho e_{t-1} + \nu_t \tag{11}$$

Further model parsimony may be achieved by imposing insignificant restrictions among the variables (Cuthbertson, Hall, and Taylor (1992)).

In Tables 5a-5d, we report our final estimated equations,obtained for each of the dollar exchange rates using this modelling strategy. In each case, the equations were estimated using data up to the end of 1988, with two years of data retained for tests of forecasting ability. In each case, the error correction exchange rate equation performs well in-sample, with well-determined coefficients. Each equation yields a highly-significant, correctly-signed coefficient on the gold price variable, easily passes a battery of diagnostic tests, and appears to display parameter stability over the remaining two years of data, as measured by the Chow and predictive failure (PF) tests (Hendry (1979)).

Note that we make no attempt to rationalize the form of the short-run dynamics of the equations reported in Tables 5a-5d. It is possible that these result from the interaction of different speeds of adjustment of wages, goods prices, and asset prices, and/or from short-term fads in foreign exchange markets (See MacDonald and Taylor (1992)). For our purposes, however, it is only necessary to note the strong significance of the gold price variable in these equations, even after controlling for complex data dynamics.

As a point of comparison, we also conducted dynamic post-sample prediction tests of the kind reported earlier. That is to say, for the twenty-four observations, 1989(1) - 1990(12), we allowed the parameters to be re-estimated at each data point (although the cointegrating parameter vector was held constant) and computed dynamic forecasts for a number of months ahead. The resulting U-statistics from this exercise are reported in Table 6. Except for the yen/dollar rate, these show a marked improvement over those reported for the VAR system in Tables 2a-2e. This effect is particularly marked for the sterling/dollar exchange rate.

IV. Concluding Remarks

The general conclusion which emerges from the various empirical investigations we have conducted is that gold price movements have explanatory power with respect to exchange rate movements, over and above the effects of movements in monetary fundamentals and other variables that enter standard exchange rate models. Based on the concept of gold as "an asset without a country" and the argument that changes in country preferences will be systematically reflected in the price of gold, our empirical findings can be interpreted as indirect evidence that exchange rate movements are largely coterminus with events that change preferences for holding claims on different countries. Further work on this issue might concentrate on developing a more complete theoretical framework capable of suggesting more rigorous, testable restrictions on the data.

Table 1. Estimates of a Single Equation Monetary Model Including and Excluding Gold and Wheat Prices 1/

Exchange Rate	$\Delta(m-m^*)$	$\Delta(y-y^*)$	$\Delta(r-r^*)$	$\Delta(\pi-\pi^*)$	Δg	Δw	R^2	D.W.
UK£/US$	-0.464	0.203	-0.158	0.449	--	--	0.03	1.86
	(-1.437)	(1.235)	(-0.617)	(1.180)				
	-0.387	0.213	-0.282	0.501	-0.176	--	0.15	1.86
	(-1.275)	(1.382)	(-1.171)	(1.405)	(-5.002)			
	-0.465	0.202	-0.158	0.450	--	-0.638E-2	0.03	1.86
	(-1.435)	(1.229)	(-0.617)	(0.381)		(-0.101)		
DM/US$	-0.104	0.142	0.572	-0.887	--	--	0.04	2.15
	(-0.409)	(1.167)	(1.964)	(-1.276)				
	0.135	0.121	0.289	0.200	-0.219	--	0.22	2.08
	(0.579)	(1.096)	(1.080)	(0.307)	(-6.288)			
	-0.103	0.142	0.572	-0.886	--	0.318E-2	0.04	2.15
	(-0.400)	(1.164)	(1.959)	(-1.269)		(0.051)		
¥/US$	-0.017	-0.145	0.738	-0.718	--	--	0.04	1.87
	(-0.070)	(-0.789)	(2.336)	(-1.659)				
	0.018	-0.109	0.561	-0.771	-0.138	--	0.12	1.80
	(0.079)	(-0.613)	(1.817)	(-1.845)	(-3.700)			
	0.014	-0.160	0.749	-0.703	--	-0.075	0.05	1.87
	(0.058)	(-0.872)	(2.371)	(-1.625)		(-1.151)		

Table 1 (cont)

Exchange Rate	$\Delta(m-m^*)$	$\Delta(y-y^*)$	$\Delta(r-r^*)$	$\Delta(\pi-\pi^*)$	Δg	Δw	R^2	D.W.
FFR/US$	0.040	0.202	0.362	-0.605	--	--	0.03	2.07
	(0.208)	(1.412)	(1.405)	(-0.944)				
	0.077	0.205	0.187	0.066	-0.201	--	0.20	2.08
	(0.435)	(1.581)	(0.792)	(0.112)	(-6.077)			
	0.039	0.200	0.367	-0.604	--	-0.018	0.03	2.08
	(0.203)	(1.398)	(1.416)	(-0.941)		(-0.299)		
DM/¥	-0.278	0.025	-0.398	-0.342	--	--	0.02	1.89
	(1.369)	(0.256)	(0.875)	(-0.877)				
	-0.181	0.020	-0.420	0.158	-0.187	--	0.19	2.01
	(-0.972)	(0.227)	(-1.010)	(0.433)	(-6.039)			
	-0.232	0.022	-0.378	-0.322	--	0.063	0.03	1.89
	(-1.118)	(0.226)	(-0.830)	(-0.824)		(1.112)		

1/ The dependent variable is the change in the logarithm of the exchange rate. Unstarred explanatory variables in the exchange rate equations for the US$ (or in the DM/¥ equations) are associated with the United States and the dollar (or Japan and the yen). In particular, m is the logarithm of the U.S. narrow money stock, y is the logarithm of U.S. industrial production, r is a short-term dollar interest rate, π is the U.S. rate of inflation over the previous 12 months, g is the logarithm of the price of gold in dollars, and w is the logarithm of the price of wheat in dollars (g and w are expressed in yen in the DM/¥ equations). Starred explanatory variables are associated with the other country/currency. R^2 and D.W. denote the coefficient of determination and Durbin-Watson statistic respectively; figures in parentheses are t-ratios.

Table 2a. Theil's U-Statistic for VAR Exchange Rate Forecasts 1/

		System Including Gold				System Excluding Gold	
		Actual Future Values of Exogenous Variables		Forecast Values of Exogenous Variables		Actual Future Values of Exogenous Variables	Forecast Values of Exogenous Variables
Exchange Rate	Months Ahead	Actual Future Gold Price	Forecast Gold Price	Actual Future Gold Price	Forecast Gold Price		
UK£/US$	1	1.42	1.42	1.42	1.42	1.27*	1.28
	2	1.59	1.58	1.38	1.42	1.23*	1.28
	3	1.67	1.66	1.24	1.30	1.14*	1.23
	4	1.63	1.63	1.03	1.01*	1.08	1.09
	5	1.60	1.62	0.92	0.86*	1.06	1.01
	6	1.61	1.66	0.87	0.80*	1.16	0.96

Table 2a(cont.)

7	1.62	1.66	0.85	0.75*	1.27	0.91
8	1.59	1.59	0.88	0.77*	1.35	0.91
9	1.67	1.62	0.88	0.74*	1.51	0.89
10	1.73	1.63	0.93	0.76*	1.59	0.89
11	1.79	1.64	0.96	0.77*	1.72	0.88
12	1.82	1.64	0.94	0.73*	1.84	0.83

1/ The VAR systems included domestic and foreign money supplies, output, interest rates, and inflation as well as the exchange rate. Theil's U-Statistic as the ratio of the root mean square error (RMSE) of the VAR forecast to the RMSE of the random walk forecast. An asterisk denotes minimum U-statistic for that forecast horizon.

as static) is highly correlated with publicly available information and the bidding behavior of firms owning neighboring leases, but it is not correlated with the bidding behavior of uninformed firms, conditional on the publicly available information. (See Hendricks and Porter (1988).)

A second strong assumption is that there is only one informed bidder on drainage leases. In fact, there are on average 3.87 neighboring leases, as indicated in Table 6. However, there are both institutional and empirical reasons to believe that the informed bidders will coordinate their actions, and effectively bid as one. There are two institutional reasons. First, joint bids are legal, as described above in Section 3. Second, tracts sharing a common pool are typically unitized, to avoid inefficiencies associated with overdrilling. (See Libecap and Wiggins (1985) for more detail.) A unitization agreement allocates revenues from a common pool according to a pre-specified scheme, typically on the basis of acreage above the pool, and serves as an institution to facilitate side payments. (In addition, there is the threat to end the unitization agreement and overdrill, should anyone break an agreement.)

The empirical reasons are several, as well. First, multiple informed bids on a tract were relatively uncommon, as indicated in Table 6. Table 6 reports the frequency distribution of the number of neighboring leases, where there is at least one adjacent lease, as well as the number of bids submitted by firms owning neighboring leases (informed bids), and by non-neighbors (uninformed bids). Note that the frequency distribution of the number of neighbor bids is almost the same before and after 1970, with mean about one, despite the increase in the average number of adjacent tracts after 1970. Second, multiple neighbor bids tended to occur on high value tracts, and ex post returns were higher than on single bid tracts, rather than lower, as might be expected from competitive bidding. Finally, the potential winner's curse problems faced by uninformed bidders are augmented by the presence of multiple competing informed bidders. If uninformed bidders have access to public signals alone, they should not participate. Yet the bidding of uninformed bidders appears to be independent of the number of firms owning neighboring leases. All three of these facts are consistent with coordinated bidding by informed firms, where multiple bids are occasionally submitted to create the appearance of competition. (See Porter and Zona (1992) for an account of a collusive scheme that similarly relied on non-serious bids.) The facts are also consistent with one of the informed bidders having superior information, and the others being akin to the uninformed firms, with the same empirical predictions.

The third and fourth predictions, concerning profits, are borne out by the data, as demonstrated in Table 7, which differentiates between tracts won by neighbors and non-neighbors, and within those categories depending on whether the other type of firm submitted a bid. Profits are reported only for the 1959-1973 subsample, where the figures are more reliable. Consistent with the third prediction above, uninformed firms break even approximately, and lose money on tracts where no informed firm bids.

On drainage tracts, HPB calculate that firms capture about a third of social rents, compared to a quarter on wildcat tracts. Nevertheless, fewer bids are submitted on average (2.45 on drainage leases versus 3.52 on wildcat leases, in the 1954-1979 sample). Entry appears to be inhibited by informational barriers to entry, and non-neighboring firms break even on average.

As for the first two predictions, as illustrated in Figure 2b, only the first is borne out, as Figure 3 demonstrates. Figure 3 depicts the empirical distribution function of the highest informed and uninformed bid submitted on the 295 drainage tracts that were offered for sale and received bids in the period 1959-1979. The informed firms indeed bid more often, as indicated by the height of the distribution functions at zero, and submit the highest bid more often (on 61.4 percent of the leases). However, there is no evidence of a mass point at the announced reserve price, which is about $62,500 (at $25 per acre for 2,500 acres, the average drainage tract size). Nor do the distribution functions coincide above the reserve price, although they are similar above $4 million. The striking aspect of Figure 3 is not that uninformed firms submit bids less often, but rather that when they bid, they tend to submit high bids.

Another assumption of the preceding theory is that the government accepts all bids above the announced reserve price. On the contrary, they rejected 58 of the 295 high bids submitted on drainage tracts, or 19.7 percent. Table 8 compares bidding on accepted and rejected drainage tracts. Two aspects are of note. First, a higher fraction of rejected bids are by informed firms. Second, the government is much more likely to reject a bid if it is low, in an absolute sense. (This is analogous to the rejection policy on wildcat tracts, as described in Section 3.)

As HPW demonstrate, it is possible to reconcile the disparities between the predictions depicted in Figure 2b and the empirical distribution of Figure 3, if one accounts for the propensity of the government to reject low bids. Consider the previous example, but now assume that there is an unannounced tract specific reserve price, unknown to the bidders, that is distributed uniformly on the interval [1, 3], where 1 is the announced minimum bid. Then a bid b between 1 and 3 will be accepted with probability (b-1)/2. Assume also that the reserve price is determined prior to the bidding, and unaffected by submitted bids. Then denote by $\beta_0(v)$ the optimal bidding strategy of the informed firm when there is no uninformed bidder. Here $\beta_0(v) = (1 + v)/2$ for v in [1, 5], and $\beta_0(v) = 3$ for $v > 5$, as depicted in Figure 4a. When there is an uninformed bidder present, the equilibrium bidding strategy of the informed bidder is $\beta_1(v) = \max\{\beta_0(v), \beta(v)\}$, as depicted by the solid line in Figure 4a. That is, for low value tracts, the informed firm is concerned with the possibility of having its bid rejected, and so increases its bid. The effect of this increase is to knock out low uninformed bids. Low uninformed bids now earn negative expected profits, because the bidding strategy of the informed firm is more aggressive than what a zero profit calculation would entail. In the Figure, β_0 lies to the right of β for bids less than 3. The implications for the bid distribution functions are shown in Figure 4b. There is no longer a mass point in the informed firm bid distribution function at the reserve price, and the uninformed firm no longer submits low bids. The distribution functions coincide above 3, the upper bound on the support of the reserve price. The rest of the predictions from the simple model remain valid. In the drainage leases, only 6 of the 122 bids above $4 million were rejected, and the empirical distributions essentially coincide above that level. Thus a simple adaptation of the theory can account for the bidding behavior on drainage leases. The fact that informed firms submit a higher percentage of rejected bids is consistent with the prediction that they are more likely to bid low, and low bids are more likely to be rejected.

The theory is too simple in that it assumes that the government has no private information of its own, and because the bidders do not account for the possibility of a reoffering in the event that the low bid is rejected. On the latter point, it is notable that less than a third of the tracts with rejected bids were reoffered, and reofferings occurred a year and a half later on average. (See Hendricks, Porter, and Spady (1989).) Therefore, it is not unreasonable to assume that firms ignore the possible repercussions of their bidding for future reofferings. On the former point, the government also has access only to seismic information before the auction, and submitted bids do not seem to influence reserve prices, except when more than three bids are submitted. (The informed firms may submit multiple bids on valuable tracts precisely to manipulate bid adequacy decisions in these cases.) As HPW demonstrate, if one accounts for private information observed by the government, then the theoretical predictions of the example above continue to hold. They show that the distribution of the informed bid should stochastically dominate that of the maximum uninformed bid, and the distributions should coincide above the support of the reserve price. These predictions are satisfied by the empirical distribution in Figure 3.

Therefore, a theoretical model that accounts for important institutional features can describe the data fairly accurately. The model emphasizes informational asymmetries, rather than cost asymmetries. While cost asymmetries are undoubtedly present, I believe that their influence is swamped by informational asymmetries. A model of cost asymmetries alone cannot account for the lack of correlation between the uninformed bids and ex post tract values. Also, cost asymmetries should be mitigated by unitization agreements, which encourage efficient production plans. In contrast, several predictions of an auction model with asymmetric information are confirmed by the bidding data, after the government rejection decision is accounted for.

6. AN ASSESSMENT OF THE MECHANISM

Is the OCS auction mechanism optimal? Essentially, the issues are: whether undue rents are being captured by the firms in the bidding game, either because of lack of competition, capital constraints, or insufficient (or asymmetric) information; whether rents are dissipated via excess drilling, due to costly duplication of effort in generating common information; and whether the rate at which tracts have been offered for sale has been sensible.

In some important respects, the OCS leasing program is well designed. Bidding for wildcat leases appears to be relatively competitive, and the government probably captures a reasonable share of the rents, given the risks involved.

Owners of adjacent leases extract sizable information rents in drainage lease sales. To the extent that profits in subsequent drainage sales are anticipated, expected future drainage rents are likely to be reflected in bidding for wildcat leases. If subsequent profits are not anticipated, perhaps because drainage sales do not always follow wildcat sales, then the government could increase royalty rates on drainage leases. The Bayesian Nash equilibrium of the asymmetric bidding game predicts that non-neighbors earn zero expected profits, if they have access only to public information. Then a higher royalty rate serves as a tax solely on the firms owning neighboring leases, and with access to superior information. The problem is that for tracts with relatively small deposits, it would no longer pay to bid at all. In addition, a higher royalty rate exacerbates the moral hazard problem of less ex post exploration than is socially optimal. These arguments assume that a royalty on revenues would be employed, given the difficulties in measuring costs. Alternatively, royalties might apply only to revenues above some prespecified estimate of likely drilling costs, based on industry averages. Nevertheless, some caution is in order, since changing the rules of drainage auctions would probably alter bidding and exploration decisions on wildcat leases, which are qualitatively much more important.

More troubling is the apparent delay of exploration decisions until the end of the lease term. The fixed lease term induces a deadline effect, which may entail suboptimal overdrilling at the end of the term. However, a fixed term also reduces purely speculative motives for acquiring, and probably not exploring, a tract.

There are potential gains from the coordination of drilling programs. There may be a concern that coordination in exploration might extend into bidding. Of course, current joint bidding arrangements are potentially collusive, as are unitization agreements, and yet they appear not to have had a detrimental impact on competition. The heterogeneity in tract values, and in perceptions of values of individual tracts, as exemplified by the variation in bidding across and within tracts, must be an obstacle to cooperation, and probably accounts for the relatively competitive outcomes. Having said that, the ban on joint bids involving two or more of the largest firms seems like a sensible policy. There is a clear potential for bidding consortia to limit competition. Further, if consortia are beneficial because they raise capital, then joint bids with industry outsiders (L&F bids) serve the same purpose, and probably enhance competition. (This argument is analogous to the notion that entry by building a new plant is socially preferable to entry via acquisition of an existing plant, as competition is stimulated.)

Another issue is whether more information could be made available prior to wildcat sales. Under current practice, firms acquire a risky prospect, and royalty schemes do not provide much insurance. In particular, they do not provide any insurance for drilling costs, since royalties apply only to revenues. As on drainage sales, royalties on wildcat leases should also only apply to revenues above a predetermined level. One reason why the government sells leases is so that it does not get into the drilling business itself. However, the theory of optimal auction design when there is noncooperative bidding suggests that the government should make public as much information as possible. If collusive bidding is a concern, then a random reservation price policy, or else higher announced reserve prices on valuable tracts, can be used.

It is clear from Table 1 that there was a fundamental policy change in the 1980s, as the rate at which tracts were offered for sale increased dramatically. This increase coincided with a fall in the real price of oil and gas, and so may have been mistimed. One might argue that any tract with positive net present value should be leased as quickly as possible, given that the U.S. is probably a price taker on world oil markets (at least with respect to offshore supply). In addition, there are clear political motives to bring revenues forward. Nevertheless, the U.S. may want to delay some lease sales. The public sector has a monopoly position on offshore oil and gas rights, and may be able to raise the price it receives by restricting supply. There may have been a problem in the 1980s, as the number of tracts offered for sale may have exceeded industry capacity to explore them. Also, with so many tracts on the market, it might be easier for firms to subdivide the OCS lands, say by geographic regions, and suppress competition. The preceding discussion is speculative, but there has been much less bidding on offshore prospects since 1983.

A final issue concerns what the Department of Interior should maximize. The optimal auction design literature, and some of the above discussion, assumes that government revenue maximization is the goal. However, another reasonable goal is the expeditious exploration and development of offshore oil and gas supplies. To that end, the possibility of profits in the bidding process encourages firms to incur presale exploration expenses, and thereby identify productive tracts for the bidding and exploratory drilling stages of the process.

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Table 1: Summary of Offshore Oil and Gas Lease Sales, 1954-1990*

		Tracts Receiving Bids		Tracts Sold				Tracts with Rejected Bids	
Period	# of Tracts Offered	#	Bids per Tract	#	Mean Acreage	Total Winning Bids	Mean Winning Bid	#	Mean High Bid
1954-1960	950	454	2.94	419	4,153	621	1.481	35	0.137
1961-1967	1,460	841	2.95	801	4,672	1,317	1.645	40	0.151
1968-1974	2,041	1,269	4.04	1,103	4,779	12,855	11.655	166	1.254
1975-1982	6,811	2,753	2.59	2,383	5,207	26,591	11.159	370	1.963
1983-1990	136,952	8,011	1.38	7,582	5,313	14,394	1.898	429	1.535
1954-1990	148,214	13,328	2.03	12,288	5,163	55,778	4.539	1,040	1.542

*Dollar figures are nominal, and in millions of dollars.

Table 2: Characteristics of Wildcat Tracts 1954–1979, by Number of Bidders*

	Number of Bidders							
	1	2	3	4	5–6	7–9	10–18	Total
No. of Tracts	902	463	255	212	264	234	180	2510
B_1	1.283	2.667	4.070	5.523	7.871	14.103	21.778	5.538
	(0.087)	(0.198)	(0.375)	(0.491)	(0.605)	(1.166)	(1.355)	(0.211)
$(B_1-B_2)/B_1$	—	0.549	0.490	0.460	0.386	0.336	0.298	0.442
	—	(0.013)	(0.017)	(0.017)	(0.016)	(0.014)	(0.015)	(0.007)
No. Sold	707	424	241	207	263	233	180	2255
(fraction)	(0.784)	(0.916)	(0.945)	(0.976)	(0.996)	(0.996)	(1.000)	(0.898)
B_1	1.495	2.756	4.170	5.624	7.898	14.160	21.778	6.071
	(0.109)	(0.214)	(0.394)	(0.510)	(0.607)	(1.170)	(1.355)	(0.232)
No. Drilled	431	315	208	176	239	210	178	1757
(fraction)	(0.610)	(0.743)	(0.863)	(0.850)	(0.909)	(0.901)	(0.989)	(0.779)
No. Productive	175	148	97	90	117	132	122	881
(fraction)	(0.406)	(0.470)	(0.466)	(0.511)	(0.490)	(0.629)	(0.685)	(0.501)
Disc. Revenues	13.497	15.509	19.451	25.063	26.244	28.845	33.382	22.507
	(2.040)	(2.108)	(2.478)	(4.105)	(3.885)	(3.331)	(5.087)	(1.263)

*B_1 denotes the highest bid on a tract, and B_2 the second highest bid. Dollar figures are in millions of 1972 dollars, and refer to means of preceding sample. (So the mean of discounted revenues is for the sample of productive tracts.) Standard errors of the sample means are displayed in parentheses, except where noted.

Table 3: Number of Bids and Rejection Decisions on Wildcat Tracts*

	Number of Bidders								Mean #	Mean
	1	2	3	4	5-6	7-9	10-18	Total	of Bids	Bid
1954-1969										
Accepted bids	339	213	106	103	126	114	55	1056	3.46	2.671
	(.819)	(.982)	(.982)	(1.0)	(1.0)	(1.0)	(1.0)	(.929)	(0.09)	(0.159)
Rejected bids	75	4	2	0	0	0	0	81	1.10	0.219
	(.181)	(.018)	(.018)	(0.0)	(0.0)	(0.0)	(0.0)	(.071)	(0.04)	(0.021)
All tracts	414	217	108	103	126	114	55	1137	3.29	2.496
									(0.08)	(0.149)
1970-1979										
Accepted bids	368	211	135	104	137	119	125	1199	4.03	9.067
	(.754)	(.858)	(.918)	(.954)	(.993)	(.992)	(1.0)	(.873)	(0.10)	(0.393)
Rejected bids	120	35	12	5	1	1	0	174	1.49	1.095
	(.246)	(.142)	(.082)	(.046)	(.007)	(.008)	(0.0)	(.127)	(0.07)	(0.112)
All tracts	488	246	147	109	138	120	125	1373	3.71	8.056
									(0.09)	(0.345)
1954-1979										
Accepted bids	707	424	241	207	263	233	180	2255	3.76	6.071
	(.784)	(.916)	(.945)	(.976)	(.996)	(.996)	(1.0)	(.898)	(0.07)	(0.232)
Rejected bids	195	39	14	5	1	1	0	255	1.36	0.816
	(.216)	(.084)	(.055)	(.024)	(.004)	(.004)	(0.0)	(.102)	(0.05)	(0.081)
All tracts	902	463	255	212	264	234	180	2510	3.52	5.538
									(0.06)	(0.211)

*The numbers in parentheses are the fraction of the total, except for the mean number of bids andthe mean high bid, where they are standard errors of the sample means. Bids are in millions of 1972 dollars.

Table 4: Wildcat Bidding by Large Firms, 1954-1979*

Firm	Solo Bids	Joint Bids: L Only	Joint Bids: L&F	# of Bids	# of Wins
A/C/G/C	1036	71	439	1546	426
SOCAL (*)	493	112	262	867	281
Amoco (SOIND) (*)	197	248	374	819	213
Shell Oil (*)	551	6	184	741	251
Kerr/Marathon/Felmont	63	341	387	791	170
LaLand/Hess/Cabot	18	268	348	634	132
Sun Oil	412	158	36	606	156
Exxon (*)	522	47	32	601	197
Union Oil of Ca.	122	185	284	591	173
Gulf Oil (*)	222	122	242	586	218
Mobil (*)	83	236	146	465	199
Texaco (*)	148	174	122	444	158

*A/C/G/C refers to ARCO/Cities/Getty/Continental. L Only bids are joint bids involving only the firms listed in this table. L&F bids are joint bids involving one or more of the firms in this table and at least one other firm. The three bidding consortia listed in the table are each counted as single bidding units. Firms indicated by (*) were prohibited from joint bids with each other in 1975.

Figure 1: Hazard Rate for Exploratory Drilling on Wildcat Tracts, 1954-1990

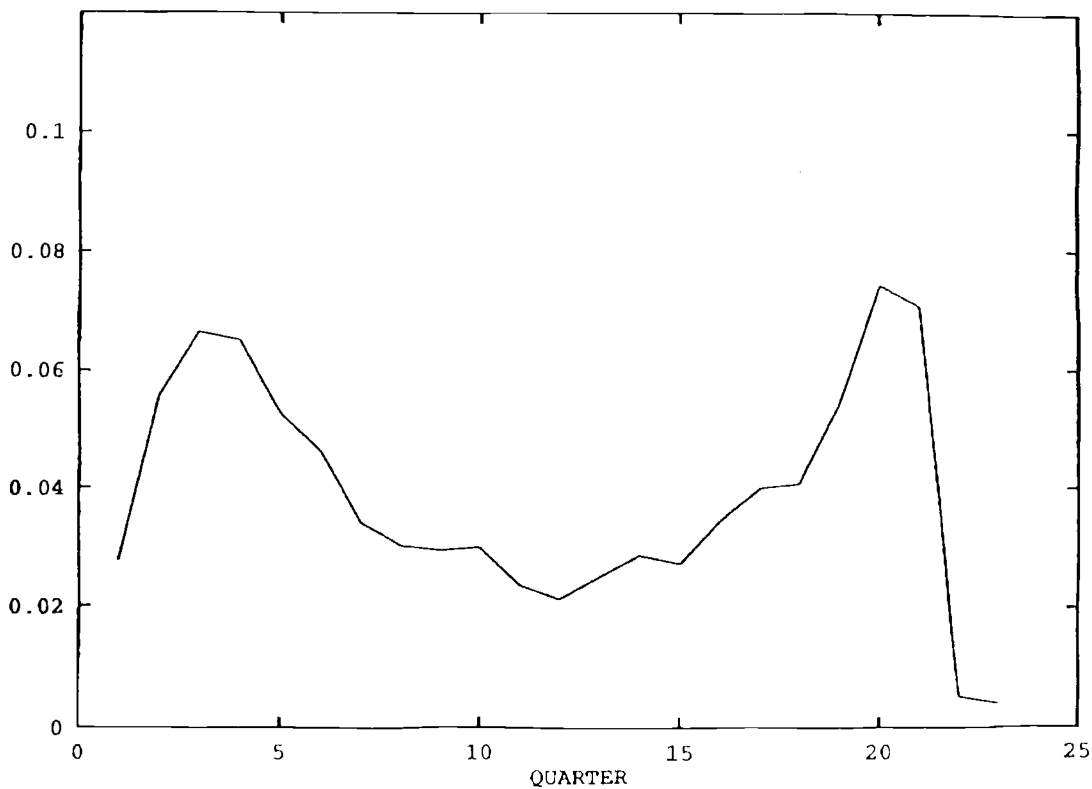


Table 5: Wildcat Tract Characteristics 1954-1979, by Year of Initial Drilling*

	Year After Acquisition				
	1	2	3	4	5
Risk Set					
Number	2184	1456	1075	879	732
BID	14.50	13.93	13.62	13.47	13.40
	(1.62)	(1.47)	(1.40)	(1.40)	(1.41)
Number of bids	3.81	2.87	2.46	2.31	2.18
	(3.28)	(2.44)	(2.04)	(1.94)	(1.81)
Tracts Drilled					
Number (fraction)	728	381	196	147	217
	(0.333)	(0.262)	(0.182)	(0.167)	(0.298)
BID	15.65	14.80	14.29	13.82	13.48
	(1.26)	(1.30)	(1.21)	(1.28)	(1.21)
BIDDIF	0.769	0.679	0.579	0.525	0.213
	(0.041)	(0.058)	(0.078)	(0.098)	(0.075)
Number of bids	5.68	4.05	3.11	2.99	2.37
	(3.90)	(3.02)	(2.35)	(2.36)	(1.81)
HIT (fraction)	413	172	91	66	85
	(0.567)	(0.451)	(0.464)	(0.449)	(0.392)
REV	16.32	15.54	15.60	15.57	15.17
	(1.50)	(1.72)	(1.72)	(1.98)	(1.53)

*Except when noted, standard deviations are displayed in parentheses. BIDDIF is the difference between the BID (the logarithm of the winning bid in 1972 dollars) and the average value of BID on tracts in the risk set that were sold in the same year. For BIDDIF, standard errors of the sample means are displayed in parentheses.

Figure 2a
Bid Function for Example with Fixed Reserve Price

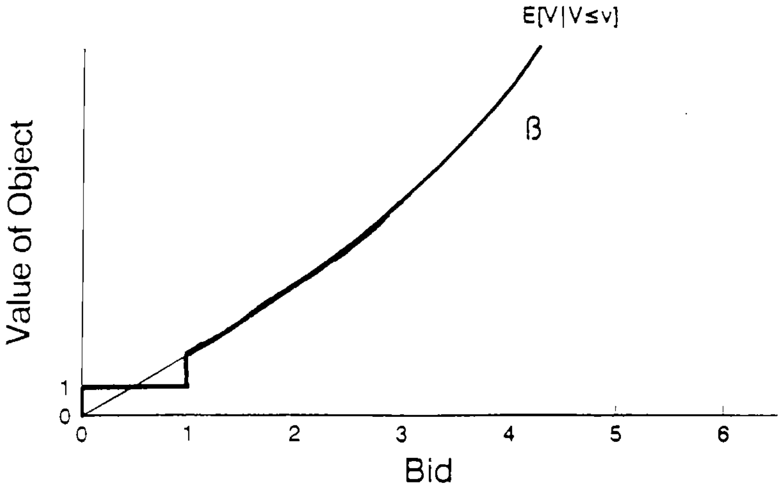


Figure 2b
Bid Distribution Functions for Example with Fixed Reserve Price

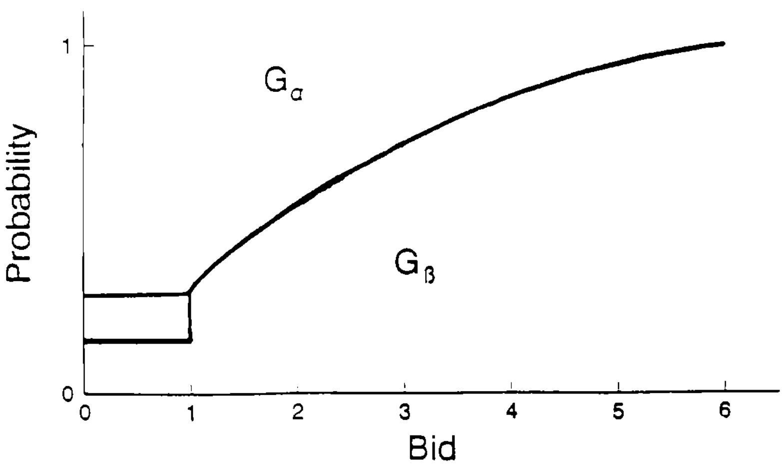


Table 6: Frequency Distributions on Drainage Tracts

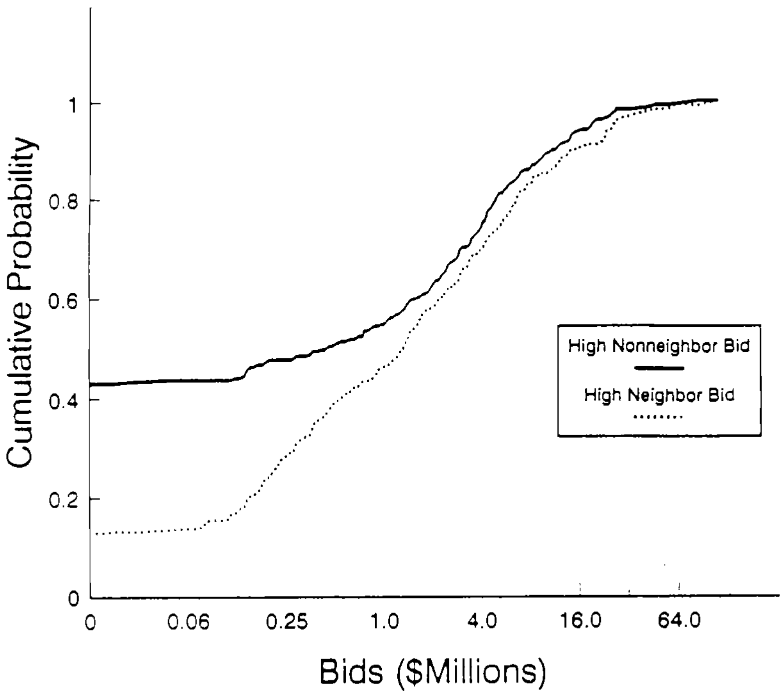
	Number							
	0	1	2	3	4	5-6	7-12	Mean
1959-1969								
Neighbor tracts	0	19	36	19	12	14	7	3.04
Neighbor bids	12	80	14	0	1	0	0	1.05
Non-neighbor bids	46	31	11	11	2	2	4	1.35
1970-1979								
Neighbor tracts	0	14	34	40	26	43	31	4.21
Neighbor bids	30	118	33	6	1	0	0	1.10
Non-neighbor bids	76	46	36	11	6	9	4	1.38
1959-1979								
Neighbor tracts	0	33	70	59	38	57	38	3.78
Neighbor bids	42	198	47	6	2	0	0	1.08
Non-neighbor bids	122	77	47	22	8	11	8	1.37

Table 7: Bidding on Drainage Tracts, by Type of Bidder[*]

	Wins by Neighbor Firms		Wins by Non-Neighbor Firms		
	No Non-Neighbor Bid	Total	No Neighbor Bid	Neighbor Bid	Total
1954-1979					
Number of tracts	77	135	32	70	102
Number drilled	60	117	27	67	94
Number productive	46	95	12	40	52
(fraction of total)	(0.60)	(0.70)	(0.38)	(0.57)	(0.51)
Mean winning bid	5.19	10.16	3.31	8.90	7.14
	(1.09)	(1.55)	(0.85)	(1.30)	(0.96)
Mean disc. revenues	11.67	19.83	4.24	18.29	13.88
	(2.60)	(3.14)	(1.57)	(4.16)	(2.96)
1954-1973					
Number of tracts	43	75	12	37	49
Mean net profits	1.56	4.93	-2.00	1.83	0.89
	(1.82)	(2.41)	(0.92)	(3.23)	(2.45)

[*]Dollar figures are in millions of 1972 dollars. Except where noted, standard errors of sample means are displayed in parentheses.

Figure 3
Distribution of Bids on Drainage Tracts, 1959-1979



All bids are represented in 1972 dollars.

Table 8: Comparison of Accepted and Rejected Drainage Bids, 1959-1979*

	Accepted	Rejected
Largest neighbor bid	7.047 (0.971)	1.165 (0.310)
Largest non-neighbor bid	5.088 (0.667)	0.319 (0.116)
High bid	8.861 (0.978)	1.453 (0.312)
Number of neighbor bids	1.11 (0.04)	0.93 (0.07)
Number of non-neighbor bids	1.62 (0.13)	0.33 (0.09)
Number of bids	2.73 (0.14)	1.26 (0.30)
Number of neighbor tracts	3.73 (0.14)	4.00 (0.30)
Fraction with high bid by neighbor	0.57	0.79
Number of tracts	237	58

*Bids are in millions of 1972 dollars. The numbers in parentheses are standard errors of the sample means.

Figure 4a
Bid Functions for Example with Unknown Reserve Price

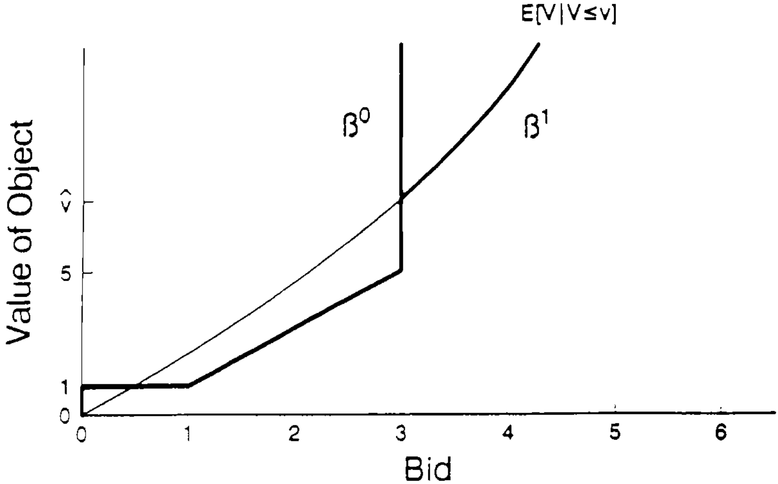


Figure 4b
Bid Distribution Functions for Example with Unknown Reserve Price