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The Effect of Options on Stock Prices: 1973 to 1995

SORIN M. SORESCU*

ABSTRACT

I show that the effect of option introductions on underlying stock prices is best described by a two-regime switching means model whose optimal switch date occurs in 1981. In accordance with previous studies, I find positive abnormal returns for options listed during 1973 to 1980. By contrast, I find negative abnormal returns for options listed in 1981 and later. Possible causes for this switch include the introduction of index options in 1982, the implementation of regulatory changes in 1981, and the possibility that options expedite the dissemination of negative information.

AN INCREASINGLY IMPORTANT ISSUE IN THE STUDY of financial markets is the extent to which options interact with their underlying stocks. Although Black and Scholes (1973) view options as redundant securities, Ross (1976) theorizes that stock prices in incomplete markets are generally affected by the introduction of new options, to the extent that these options span additional states of nature in the Arrow–Debreu uncertainty space. More recently, Detemple and Selden (1991) and Detemple (1990) also demonstrate that general equilibrium price effects are possible when nonredundant stock options begin trading in incomplete markets. These theoretical models, however, provide little indication about the direction (or magnitude) of such price effects. (Factors such as investor endowments and risk preferences or the degree of market completeness determine how each individual stock will react at the time of its own option introduction.) Thus, theory suggests that options could raise underlying stock prices, lower them, or leave them unaffected (if general equilibrium effects are sufficiently small to have no economic significance). This theoretical ambiguity justifies the need for extensive empirical research.

* Assistant Professor of Finance, University of Houston. This paper is based on my Ph.D. dissertation at the University of Florida. I am indebted to Mark Flannery for many patient and helpful discussions. I owe special thanks to Bartley Danielsen, René Stulz, and an anonymous referee. Useful suggestions were provided by David T. Brown, David Ikenberry, Praveen Kumar, Bong-Soo Lee, Mahendrarajah Nimalendran, Jay Ritter, Subu Venkataraman, and Arthur Warga, as well as seminar participants at the University of Houston, the Federal Reserve Bank of Chicago, Rutgers University, the University of Wisconsin–Milwaukee, Michigan State University, Hofstra University, Concordia University, and York University. The dataset in my dissertation consists of options listed during 1973 to 1992. Information about options listed from 1993 to 1995 was generously provided by Bartley Danielsen. Financial support from the Social Sciences and Humanities Research Council of Canada is greatly appreciated.

Two prior studies have evaluated the effect of option introductions on underlying stock prices. Conrad (1989) examines the effect of option listings on stock prices and volatilities, using options introduced between 1974 and 1980. She finds significantly positive stock abnormal returns around option listing dates (two percent), and concludes that "options are not completely redundant securities." She attributes the positive price effect to an increase in the option dealer's demand for the underlying stock.

Detemple and Jorion (1990) also examine the effect of option introductions on stock prices for a sample of 300 stock options listed between 1973 and 1986. They find significant price increases (two percent) for the 201 options listed between 1973 and 1982, and no significant effects for 99 options listed between 1982 and 1986. They conclude that their results are consistent with the market completion hypothesis: "Contrary to the premises of classic arbitrage pricing models, option markets have a real effect on equilibrium prices and allocations. We find evidence of significant price increases in the optioned stock around listing dates." The authors attribute the absence of positive price effects during 1982 to 1986 to the introduction of options on S&P 500 futures in April 1982, which would have essentially completed the markets. Detemple and Jorion's market completion explanation, however, is based on a relatively small sample. With more than two decades of options trading history and almost 2,000 listing events now available, a more extensive study of the price effect of option introductions is warranted.

My findings in this paper are both interesting and puzzling. Although I confirm earlier research for the 1970s, I find that after 1980 the impact of options introductions on the underlying stocks is significantly negative. Although I can offer no definite explanation for this change, it is very robust to alternative methodologies.

This paper is organized as follows: Section I examines the price effect of initial options listings during 1973 to 1995, and shows that the positive price effect documented for the early years of options trading becomes negative in later years. Possible causes for this switch are examined in Section II. Section III concludes the paper with a discussion of plausible areas for further research.

I. The Price Effect of Option Introductions Revisited: 1973 to 1995

In this section I replicate and extend the previous studies by Conrad (1989) and Detemple and Jorion (1990) to include all options listed on organized exchanges during 1973 to 1995, and show that the price effect of option introductions becomes significantly negative beginning in 1981. I also show that this switch is unrelated to temporal changes in the characteristics of options and their underlying stocks.

Initial listing dates for all options traded on U.S. exchanges during 1973 to 1995 are obtained from the Chicago Board of Options Exchange. Options listed on a second exchange after they had first been trading on another exchange are not considered new listings, and are excluded from this study. Individual options are matched with the CRSP data manually, using company names and ticker symbols. When a CBOE option does not have a corresponding name in the CRSP database, the observation is dropped from the dataset. This results in 2,051 listing events, which are categorized as (1) call-only listings, (2) joint put/call listings, or (3) put-only listings. Put-only listings always occur after call options have been previously listed on the same stock. Since 1988, puts and calls are almost always listed simultaneously.

Throughout the paper the *benchmark measure* for the price effect of option introductions is the cross-sectional average of the cumulative abnormal returns (CARs) computed, as in Detemple and Jorion (1990), during the 11-trading-day window surrounding each listing event [L − 5 to L + 5].[1] Implicit in this methodology is the assumption that each listing event is *independent* from the others. CARs are calculated using the Brown and Warner (1985) market model (with value-weighted index), estimated over the 95 trading days preceding the event window [L − 100 to L − 6].

A. Descriptive Statistics

Table I describes the U.S. stock options market for each year from 1973 to 1995. The first three columns of Panel A classify the type of option listings into call-only, joint put/calls, and put-only contracts. Since stand-alone puts are *always* listed after calls, put-only events never correspond to new option listings.[2] The total number of new contracts, consisting of only call and put/call events, is shown in the panel's fourth column: Prior to 1988, the vast majority of these contracts were stand-alone calls; beginning in 1988 joint put-calls became most predominant.

The next six columns of Panel A contain an annual count of new stock options listed on each of the six standardized option exchanges. No noticeable time pattern is distinguishable for exchange listings, except for the Midwest Stock Exchange which merged with the CBOE in 1977, and the NYSE which was not permitted to introduce options until 1986. The last two columns of the panel show that multiple listings were not officially permitted until 1985, although the SEC tolerated a few multiple listings in the late 1970s.

The first four columns of Panel B show the annual mean and median underlying firm size as measured by its market value of equity. Underlying firm SIZEs are substantially larger in 1973–1974 than in the rest of the period. The next two columns examine the mean and median number of calendar days that—at the time of option introduction—have elapsed since

[1] Since neither Conrad (1989) nor Detemple and Jorion (1990) find any price effects associated with the announcement date of new option listings, I only study listing date effects in this paper.

[2] When I examine the set of put-only events, I find no significant CARs. The paper's analysis is therefore based exclusively on call-only and joint put/call listings.

Table I
Descriptive Statistics

Panel A classifies the type of new option listings into call-only (C), joint put/calls (P/C), and put-only (P) contracts, for the 1973 to 1995 period. The total number of *new* option contracts (calls and put/calls) is shown in the fourth column (C and P/C). Exchanges studied are the Chicago Board of Options Exchange (CB), American Stock Exchange (AM), New York Stock Exchange (NY), Pacific Stock Exchange (PS), Philadelphia Stock Exchange (PH), and Midwest Stock Exchange (MS). The total number of listings across exchanges may be higher than the number in the fourth column due to multiple exchange listings. Panel B provides descriptive statistics of newly optioned underlying stocks. The first four columns show market capitalization figures in both current and constant (1983) dollars (SIZE). The number of days that, at the time of option introduction, have elapsed since the underlying firm first became publicly traded (AGE) is shown in the fifth and sixth columns. The last two columns classify the underlying firms according to the exchange where they are trading when options are first listed: The New York and American Stock Exchanges form one category (NY–AM), while the National Association of Securities Dealers Automated Quote forms the other (Nasdaq). Note that no options were listed in 1979 due to a moratorium imposed by the SEC.

Panel A: Descriptive Statistics of New Option Contracts: 1973–1995

	Number of Options Listed (by type)				Number of Options Listed (by exchange)—Calls and Put/Calls Only						Contracts Originally Listed on Single vs Multiple Exchanges—Calls and Put/Calls Only	
Year	C	P/C	P	C and P/C	CB	AM	NY	PS	PH	MS	Single Exchange Listing	Multiple Exchange Listing
1973	28	—	—	28	28	—	—	—	—	—	28	—
1974	6	—	—	6	6	—	—	—	—	—	6	—
1975	91	—	—	91	36	43	—	—	12	—	91	—
1976	63	—	—	63	6	15	—	15	18	11	61	2
1977	20	—	21	20	2	4	—	9	5	5	16	4
1978	5	—	—	5	2	—	—	1	2	—	5	—
1979	—	—	—	—	—	—	—	—	—	—	—	—
1980	42	4	28	46	5	20	—	12	8	—	46	—
1981	9	3	48	12	3	3	—	3	3	—	12	—
1982	87	—	—	87	24	20	—	22	21	—	87	—
1983	30	—	—	30	7	12	—	5	7	—	31	—
1984	12	—	—	12	5	5	—	—	2	—	12	—
1985	56	—	—	56	19	23	8	13	6	—	48	8
1986	43	—	1	43	11	14	4	8	7	—	42	1
1987	114	2	—	116	22	27	16	28	26	—	113	3
1988	12	85	3	97	27	23	13	17	21	—	92	5
1989	6	89	—	95	22	27	10	19	19	—	92	3
1990	—	129	—	129	41	30	19	25	35	—	110	19
1991	19	133	—	152	49	50	34	51	50	—	98	54
1992	1	149	1	150	87	45	13	24	21	—	119	31
1993	—	236	1	236	142	70	41	41	37	—	173	63
1994	—	220	—	220	118	84	17	54	48	—	140	80
1995	—	254	—	254	134	115	18	69	60	—	151	103

Panel B: Descriptive Statistics of Newly Optioned Underlying Stocks: 1973–1995 (calls and put/calls only)

Year	Market Capitalization in Current Million Dollars		Market Capitalization in Constant (1983) Million Dollars (SIZE)		Underlying Firm AGE in Calendar Days		Exchange where the Underlying Stock Trades at Option Listing Date	
	Mean	Median	Mean	Median	Mean	Median	NY–AM	Nasdaq
1973	5703	2800	12813	6300	10052	8095	28	—
1974	4019	3248	7743	6259	14287	17351	6	—
1975	1650	1039	3106	1942	11858	14261	91	—
1976	904	611	1589	1094	10096	10058	63	—
1977	698	499	1170	843	6795	5305	20	—
1978	568	388	853	588	6713	3918	5	—
1979	—	—	—	—	—	—	—	—
1980	1175	594	1421	720	9004	6531	46	—
1981	756	732	824	797	3860	3264	12	—
1982	832	520	868	547	7308	5035	86	1
1983	700	495	700	495	6073	4260	31	—
1984	2557	1255	2447	1218	7014	5594	12	—
1985	2540	1718	2345	1626	5936	4604	33	23
1986	1667	1505	1521	1368	7123	5016	27	16
1987	2008	1617	1774	1420	7968	6398	94	22
1988	1227	1126	1024	950	6333	5758	56	41
1989	965	739	772	595	5378	4696	63	32
1990	1378	852	1051	653	6091	6358	86	43
1991	994	656	728	484	3216	1969	71	81
1992	681	480	485	347	2524	1178	67	83
1993	858	535	593	374	2747	1149	113	123
1994	1186	672	794	445	2254	1030	98	122
1995	1099	567	706	370	2973	1292	99	155

the initial listing of the firm's stock on an organized exchange (AGE).[3] One distinguishable pattern is a decrease in the value of AGE through time. This is especially true during 1993 to 1995, when almost half of the underlying stocks have been publicly traded for fewer than three years at the time of option initiation.

The last two columns of Panel B show that underlying stocks during 1973 to 1984 trade almost exclusively on the NYSE or AMEX. Nasdaq-traded stocks did not generally become optioned until 1985.

B. Two Distinct Subperiods: 1973 to 1980 versus 1981 to 1995

I first compute CARs by grouping all sample firms according to the calendar year of their first option introduction, and present the results in Table II. The leftmost section of the table presents mean and median CARs for the benchmark case in which 11-day CARs are averaged cross-sectionally, assuming that each listing event is independent from the others. Like Conrad (1989) and Detemple and Jorion (1990) I find positive price effects for options listed from 1973 to 1980. Annual mean CARs range from 1.06 percent to 3.37 percent, with t-statistics significant at conventional levels in four of the eight years. By contrast, the mean CARs become predominantly negative after 1980, and attain statistical significance in seven of the 15 years from 1981 to 1995. Results based on the median measure indicate a similar switching pattern, although the significance levels are somewhat lower.

To assess the robustness of this conclusion, I replicate the analysis for shorter event windows, and present the results for the shortest [L − 1 to L + 1] window in the table's second section. The data generally convey the same switching pattern, although the switch does not seem to occur until 1982, and its abruptness appears to be somewhat weaker. Results for five-, seven-, and nine-day event windows are qualitatively similar.

When several options are listed on the same day, their CARs are correlated, biasing upward the test statistics in the first two sections. Therefore, following Conrad (1989) and Detemple and Jorion (1990), I form equally weighted portfolios of stocks with identical listing dates and treat them as single securities. I refer to this grouping method as the *day-of-listing* method. This reduces the number of independent observations to 877. One potential problem with the day-of-listing method is that many of the 11-day windows do overlap, perhaps inducing correlation to their abnormal returns. As a remedy, I form equally weighted portfolios of stocks with overlapping 11-day windows and treat them as single securities. I refer to this as *window-*

[3] There are 93 underlying firms already listed on the CRSP monthly tapes on December 31, 1925 (the first day of the CRSP monthly database). For these firms I use the December 1925 date for the purpose of age calculations.

Table II
Cumulative Abnormal Returns by Listing Year: 1973–1995

Underlying stock CARs are computed using the Brown and Warner (1985) methodology, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. The left half of the table shows mean and median measures when CARs are averaged cross-sectionally, assuming each listing event to be independent from the others. In the right half, individual securities are first grouped into portfolios, which are then treated as single securities. Two different criteria are used in forming these portfolios: *day-of-listing*, which groups together securities with identical listing dates, and *window-overlap*, which groups together securities with overlapping 11-day event windows. The significance level of the mean measure is established using the usual t-statistic. The significance of the median measure is established using the Wilcoxon sign rank statistic. The number of observations corresponding to each method (N) is shown after the mean and median measures. Note that no options were introduced in 1979 due to a moratorium imposed by the SEC.

	Individual Listings						Portfolio Grouping (event window = 11 days)					
	Event Window = 11 days			Event Window = 3 days			Day-of-Listing Groups			Window-Overlap Groups		
Year	Mean (%)	Median (%)	N	Mean (%)	Median (%)	N	Mean (%)	Median (%)	N	Mean (%)	Median (%)	N
1973	2.13**	2.90Δ	27	0.95Δ	0.68**	27	0.91	1.07	4	0.91	1.07	4
1974	1.06	2.18	6	2.02**	1.52	6	1.06	1.06	2	1.06	1.06	1
1975	3.37***	2.94***	91	1.00**	0.24	91	1.72	2.18	21	2.51	2.18**	9
1976	1.84*	0.69	63	0.74Δ	0.52*	63	2.94**	1.07	29	4.32Δ	1.20	11
1977	1.29	0.08	20	−0.47	−0.63	20	1.18	−1.55	15	0.26	0.73	4
1978	3.18	3.54	5	0.40	1.04	5	3.80	2.11	4	3.24Δ	3.54	3
1979	—	—	—	—	—	—	—	—	—	—	—	—
1980	2.04**	1.94	46	0.91Δ	0.27	46	3.57**	3.48**	15	2.63*	3.33Δ	10
1981	−1.26	−2.41	12	1.43	0.45	12	0.83	−2.58	7	−0.06	−1.66	3
1982	−1.41*	−0.47	84	−0.99**	−0.54	84	−1.90Δ	−1.02	24	−4.62**	−2.94*	11
1983	−2.81*	−0.30	30	−0.29	−0.80	30	−3.44Δ	−3.57	13	−2.52	−2.33	8
1984	−1.06	−3.32	12	−0.54	−0.70	12	−0.54	−3.36	10	−0.53	−3.36	8
1985	−1.74	−1.49**	56	0.22	0.04	56	0.28	−0.55	21	0.99	−0.19	12
1986	0.10	−0.26	42	−0.25	−0.78*	42	0.21	−1.41	32	0.46	−0.16	13
1987	−1.25	−1.13	116	0.23	0.01	116	−2.30Δ	−0.63	50	−3.64	−1.34	10
1988	0.51	−0.24	96	0.29	−0.29	96	−0.19	−0.82	48	−0.37	−0.26	15
1989	−1.45**	−0.99	94	−0.85***	−0.88***	94	−2.12**	−1.96**	46	−2.46**	−2.41**	13
1990	0.51	0.45	128	−0.01	−0.16	128	0.92	−0.27	63	0.48	0.06	12
1991	−1.88***	−1.63Δ	150	−0.30	−0.13	150	−2.64**	−2.55**	62	−1.81**	−1.22Δ	10
1992	−1.63	−0.92	149	0.04	−0.19	149	−0.92	0.43	81	−2.38**	−2.71**	14
1993	−3.15***	−2.45***	234	−0.87**	−0.48Δ	234	−3.41***	−2.39***	123	−3.04***	−2.02**	15
1994	−2.04***	−1.27***	219	−0.69**	−0.66*	219	−1.99**	−1.36**	107	−2.95**	−1.98**	14
1995	−1.19*	−1.34**	244	0.35	−0.34	244	−0.84	−2.06**	100	−0.79	−1.51	15

***, **, *, Δ Indicate statistical difference from zero at the 1, 5, 10, and 15 percent levels, respectively.

overlap grouping.[4] This method further reduces the number of observations to 214. Results using both portfolio methods are presented in the third and fourth sections of Table II, where the same switching pattern emerges. The abruptness of the switch is even better illustrated in Figure 1, which depicts the time pattern of mean and median 11-day CARs, based on individual securities (Panel A), and the day-of-listing portfolio grouping method (Panel B).

I continue my analysis with a detailed examination of daily abnormal returns inside the event window. In view of the apparent switch illustrated in Table II, I divide my sample into a pre- and post-switch period using the end of 1980 as a cutoff date, and present the results in Table III. Panel A shows daily and cumulative abnormal returns inside the 11-day window for options listed from 1973 to 1980. Along the vertical axis the panel is divided into two sections, containing mean and median abnormal returns, along with corresponding test statistics. Along the horizontal axis the panel is divided into three sections, each representing a different grouping method. Overall, the panel presents *six* measures of daily abnormal returns for each trading day inside the window. Throughout the panel, most daily abnormal returns are positive. Moreover, listing date abnormal returns are always significantly positive. These findings are consistent with the results obtained by Conrad (1989) and Detemple and Jorion (1990) during that same period.

Panel B of Table III repeats the analysis for options listed from 1981 to 1995. In contrast to Panel A, daily abnormal returns during this later period are generally negative, and attain significance (15 percent level) approximately two-thirds of the time. Somewhat puzzling are the listing date abnormal returns, which are only significant in the case of median measures, and only for individual listings and day-of-listing portfolio grouping. Mean abnormal returns are negative (although not significant) only in the case of individual listings. By contrast, both mean and median abnormal returns on the day *following* the listing date are always significantly negative at better that the 5 percent level. Overall, the cumulative 11-day effect of option introduction is approximately -1.50 percent, and is always significantly negative.

[4] One problem with the *window-overlap* method is the presence of multiple, sequential overlaps, especially toward the end of my sample period. For example, during 1994 and 1995, options were listed almost continuously, with approximately five to 10 trading days elapsing between adjacent listings. A strict application of the *window-overlap* grouping method would have resulted in averaging the price effect of all 474 listings during 1994–1995 and treating that average as a single security. A more sensible application of the *window-overlap* method is to impose an arbitrarily chosen maximum time lapse between the first and the last listing grouped under the same observation. I use 15 trading days for such maximum. That is, when multiple overlaps would have caused listings more than 15 trading days apart to be grouped under the same observation, only listings occurring during the first 15 days are treated as a single observation. Listings occurring in the 16th day or later are treated as a separate observation.

The general conclusion from Table III is not altered when I use longer benchmark estimation periods, equally weighted returns (rather than value-weighted), or a constant returns model in which event-window and pre-event raw returns are compared (not shown).[5]

C. A "Switching Means" Model

Although Table II suggests the presence of a regime switch around 1981, a more careful test of this assertion is warranted. A plausible alternative is to assume that the price effect follows a time invariant distribution with negative mean throughout the sample period. Formally, I test the following two hypotheses:

[H0]: *The price effect of options is fundamentally negative throughout the entire 1973 to 1995 sample period. The observed price effect is a random variable, which follows a time-invariant distribution having a negative mean.* (Under this null hypothesis, the positive results observed during 1973 to 1980 are simple outliers, and can be explained by the fact that there is always a reasonable probability of observing repeated positive values of a random variable drawn from a negative-mean distribution.)

[H1]: *A switch in the mean CAR occurs around 1981, causing the price effect to be fundamentally positive in the pre-switch period, and fundamentally negative thereafter. This is equivalent to the price effect following a positive-mean distribution before the switch date, and a negative-mean distribution after that date.*

To discriminate between the two hypotheses, I estimate a switching means model, adapted from Kon and Jen (1978):

$$\mathrm{CAR}_i = I_i\alpha_1 + (1 - I_i)\alpha_2 + \varepsilon_i, \tag{1}$$

[5] I also form a subsample of 203 stocks listed between 1987 and 1992, for which bid-ask data are available on the Nasdaq–NMS tape. For this subsample, I recalculate CARs using bid-to-bid prices, to determine whether the negative CARs computed from transaction data reflect real price changes, or whether they merely result from more transactions occurring at the bid price during the event window. If the negative CARs illustrated in Panel B of Table III reflect real price changes, CARs should remain negative when bid-to-bid prices are used instead of transaction prices. As expected, I find significantly negative CARs using this method (-1.62 percent, $t = -2.26$) and conclude that the post-1980 negative price effect is not likely the result of an unusually large number of transactions occurring at the bid price inside the listing window.

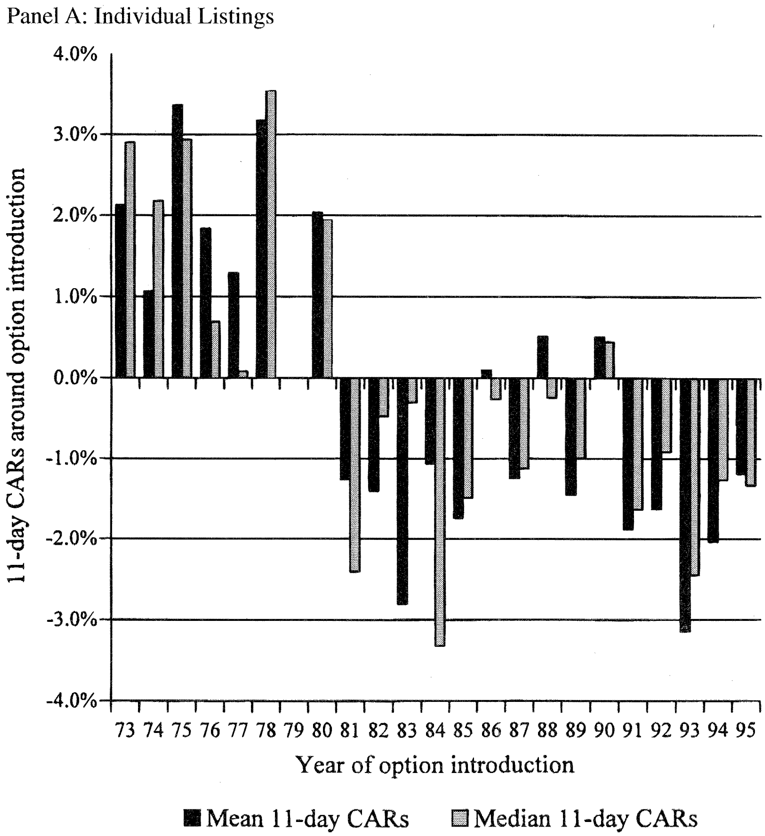

Figure 1. Eleven-day cumulative abnormal returns of underlying stocks around option introduction dates by listing year: 1973–1995. Underlying stock CARs are computed using the Brown and Warner (1985) methodology, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. Panel A shows mean and median measures when CARs are averaged cross-sectionally, assuming each listing event to be independent from the others (individual listings). In Panel B individual securities are first grouped into portfolios according to the date of their option introduction. Following this grouping, each portfolio is treated as a single security (day-of-listing portfolio grouping). Note that no options were listed in 1979, due to an SEC-imposed moratorium.

where i denotes an individual listing event, $i = 1 \ldots N$; $I_i = 1$ if $t(i) \leq T$ and $I_i = 0$ if $t(i) > T$; $t(i)$ = time of individual listing event; and T = switching date.

The three parameters to be estimated in equation (1) are the pre-switch mean CAR (α_1), the post-switch mean CAR (α_2), and the optimal switch date T. Since equation (1) is nonlinear only in T, these parameters may be estimated using a grid-search technique consisting of estimating a separate OLS

Panel B: Day-of-Listing Portfolio Grouping

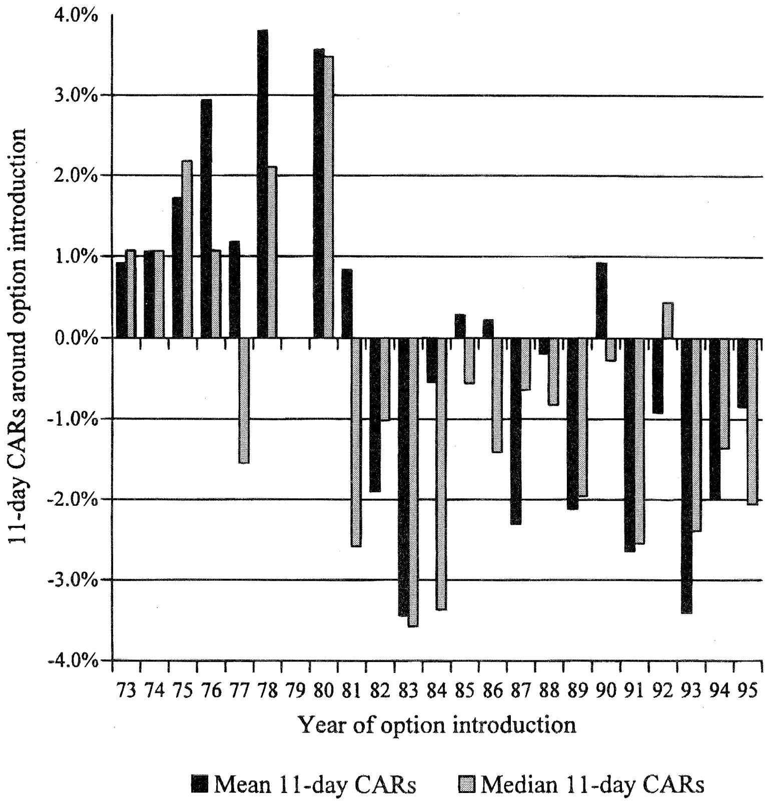

Figure 1. Continued

version of equation (1) for all possible T candidates.[6] The value of T that maximizes the likelihood function represents the optimal switching date, and the corresponding parameters α_1 and α_2 provide estimates of the pre- and post-switch means respectively. The optimal switching means model (1) can be compared to the single mean model

$$\text{CAR}_i = \alpha_0 + \varepsilon_i \tag{2}$$

using a likelihood ratio test, where α_0 represents the value of the single mean. Model (2) may be viewed as a constrained version of model (1), in which T is set equal to $t(N)$. Therefore, if model (1) provides a significantly better fit

[6] The universe of T candidates consists of $\{T = t(i),\ i = 11 \text{ to } N - 10\}$, where the 20 extreme observations are excluded, as suggested in Kon and Jen (1978).

Table III
Cumulative Abnormal Returns for the 11-Day Window around Option Listings for the Pre- and Post-switch Periods

Abnormal returns are computed using the Brown and Warner (1985) method, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. The table reports a daily abnormal return measure (DAR) as well as a cumulative abnormal return measure computed from day L − 5 (CAR). Panel A shows data for 1973 to 1980, when most abnormal returns are positive. Panel B shows data for 1981 to 1995 when abnormal returns are generally negative. In the first two columns of each panel, abnormal returns are averaged cross-sectionally, assigning equal weight to each individual listing event. In the last four columns individual securities are first grouped into portfolios, which are then treated as single securities. Two different criteria are used in forming these portfolios: *day-of-listing*, which groups together securities with identical listing dates, and *window-overlap*, which groups together securities with overlapping 11-day event windows. The number of observations corresponding to each grouping method (N) is shown at the top of each column. The significance level of the mean measure is established using the usual t-statistic. The significance of the median measure is established using the Wilcoxon sign rank statistic.

Panel A: Pre-switch Period (1973–1980)

	Days around Listing Date	Individual Listings (N = 258)		Day-of-Listings Portfolios (N = 90)		Window-Overlap Portfolios (N = 42)	
		DAR (%)	CAR (%)	DAR (%)	CAR (%)	DAR (%)	CAR (%)
Mean	L − 5	0.00	0.00	0.02	0.02	−0.18	−0.18
	L − 4	0.19	0.19	0.19	0.21	0.34*	0.16
	L − 3	0.43***	0.62***	0.23	0.44	0.39	0.55
	L − 2	0.12	0.74***	0.21	0.64*	0.09	0.64 Δ
	L − 1	−0.01	0.73***	0.19	0.83*	0.19	0.83 Δ
	L	0.64***	1.37***	0.46***	1.30***	0.58***	1.40**
	L + 1	0.23*	1.60***	0.23	1.53***	0.39	1.80**
	L + 2	0.10	1.70***	0.28*	1.81***	0.23	2.02***
	L + 3	0.19*	1.89***	0.22	2.03***	0.12	2.14***
	L + 4	0.25**	2.14***	0.28	2.31***	0.47	2.61***
	L + 5	0.33**	2.47***	0.06	2.37***	0.05	2.66***
Median	L − 5	−0.11	−0.11	0.02	0.02	0.03	0.03
	L − 4	0.14	0.22 Δ	0.05	0.29	0.27	0.16
	L − 3	0.25*	0.46 Δ	0.24 Δ	0.37	0.40	0.67**
	L − 2	−0.04	0.43 Δ	0.12	0.25 Δ	0.17	0.46
	L − 1	−0.20*	0.54*	−0.11	0.40 Δ	−0.14	0.55
	L	0.46***	0.96**	0.46***	1.24**	0.69***	1.17***
	L + 1	0.10	1.02**	0.05	0.81 Δ	0.03	0.80***
	L + 2	0.14	1.11***	0.04	0.95***	0.10	1.17***
	L + 3	−0.03	1.31**	0.11	1.56**	0.13	1.70***
	L + 4	0.16*	1.37***	0.11	1.60**	0.19*	2.08***
	L + 5	0.17	1.79***	0.01	1.27**	−0.08	1.59***

Panel B: Post-switch Period (1981–1995)

	Days around Listing Date	Individual Listings (N = 1666)		Day-of-Listings Portfolios (N = 787)		Window-Overlap Portfolios (N = 173)	
		DAR (%)	CAR (%)	DAR (%)	CAR (%)	DAR (%)	CAR (%)
Mean	L − 5	−0.13*	−0.13*	−0.12	−0.12	−0.21**	−0.21**
	L − 4	−0.17**	−0.29***	−0.16 Δ	−0.28**	−0.23**	−0.44***
	L − 3	−0.24***	−0.54***	−0.32***	−0.59***	−0.12	−0.56***
	L − 2	−0.08	−0.62***	0.00	−0.59***	−0.12	−0.68***
	L − 1	−0.09	−0.71***	−0.08	−0.66***	−0.12	−0.80***
	L	−0.01	−0.72***	0.07	−0.60**	0.02	−0.77***
	L + 1	−0.16**	−0.88***	−0.25***	−0.84***	−0.22**	−0.99***
	L + 2	−0.11*	−0.99***	−0.18**	−1.03***	−0.19*	−1.18***
	L + 3	−0.16**	−1.15***	−0.10	−1.13***	−0.22**	−1.40***
	L + 4	−0.10	−1.25***	−0.09	−1.22***	−0.05	−1.45***
	L + 5	−0.26***	−1.51***	−0.30***	−1.52***	−0.14	−1.59***
Median	L − 5	−0.17***	−0.17***	−0.21***	−0.21***	−0.16**	−0.16**
	L − 4	−0.22***	−0.36***	−0.26**	−0.33***	−0.21***	−0.43***
	L − 3	−0.15***	−0.59***	−0.18**	−0.63***	−0.13	−0.62***
	L − 2	−0.09	−0.67***	−0.03	−0.52***	−0.13	−0.85***
	L − 1	−0.05	−0.58***	−0.02	−0.55**	−0.07	−0.80***
	L	−0.10**	−0.65***	−0.12*	−0.46**	−0.06	−0.84***
	L + 1	−0.23***	−0.78***	−0.26***	−0.97***	−0.25***	−1.26***
	L + 2	−0.11**	−0.80***	−0.12 Δ	−1.06***	−0.14***	−1.33***
	L + 3	−0.19***	−1.01***	−0.12*	−1.23***	−0.13*	−1.25***
	L + 4	−0.15***	−1.00***	−0.16*	−1.26***	−0.23**	−1.27***
	L + 5	−0.25***	−1.11***	−0.26***	−1.35***	−0.14	−1.59***

***, **, *, Δ Indicate statistical difference from zero at the 1, 5, 10, and 15 percent levels, respectively.

for the data, the value of its log likelihood function should be significantly higher than that of model (2). Twice the difference between the log likelihood of model (1) and the log likelihood of model (2)—a statistic that I call "λ" below—is asymptotically distributed chi-square with one degree of freedom.[7]

Table IV presents the results. The table's top section provides estimates for the single mean model of equation (2), the middle section presents estimates of equation (1)'s switching means model, and the bottom section provides the values of the λ statistic comparing the two models. (A significant value of λ indicates rejection of H0 in favor of H1.) The five columns of Table IV correspond to alternative grouping methods and event windows. The first column provides results for the benchmark case. Under H0, CARs follow a single distribution with mean -0.971 percent, which is significantly negative ($t = -4.44$). Under H1, CARs follow a positive-mean distribution prior to some optimal switching date, after which they follow a negative-mean distribution. The optimal switching date, which maximizes the model's likelihood function, occurs between December 12, 1980 and June 29, 1981. Using a likelihood ratio test, it can be shown that a 95 percent confidence interval around this date extends from August 8, 1980 to January 26, 1982.[8] The mean CAR for options listed prior to the optimal switch date is significantly positive (2.47 percent, $t = 5.31$), which is consistent with previous studies. By contrast, the mean CAR becomes significantly negative for options listed after the switch, averaging -1.51 percent ($t = -6.29$). The difference between the two means is also statistically significant ($t = -7.60$). The likelihood ratio test shows that the constant-mean model (H0) is rejected in favor of H1 at better than the 0.1 percent confidence level. I repeat the analysis with CARs computed over nine-, seven-, five-, and three-day event windows. In all instances, the data reject H0 in favor of H1, and the optimal switch date remains in the neighborhood of 1981. Additionally, the mean difference between the two subgroups remains statistically significant at the 1 percent level. Illustrative results for the shorter three-day window are provided in the table's second column.

I repeat the analysis by grouping CARs into day-of-listing and window-overlap portfolios and varying the length of the event window in each case. The third column in Table IV shows that the results obtained with 11-day CARs computed using *day-of-listing* portfolios closely resemble the benchmark case. The results also remain qualitatively identical for the nine-, seven-,

[7] An implicit assumption in this procedure is that successive values of CAR through time are not serially correlated. That is, except for the mean switch occurring at time T, CAR$(t + 1)$ is independent from CAR(t), for all values of t. I test this assumption using a simple AR1 model and find that the autocorrelation coefficient for the entire sample is close to zero and insignificant ($\rho = 0.01073$, $t = 0.46$).

[8] Let T^* be the optimal switch date for equation (1), and $L(T)$ be the value of equation (1)'s likelihood function estimated at any given value of T. The boundaries of the 95 percent confidence interval are the two switching dates $T1$ and $T2$ such that (i) $T1 < T^*$, (ii) $T2 > T^*$, and (iii) $L(T1)$ and $L(T2)$ are both statistically different from $L(T^*)$ at the 2.5 percent significance level.

Table IV

A Formal Switching Means Model

The dependent variable is the cumulative abnormal return (CAR), computed using the Brown and Warner (1985) methodology, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. The dependent variable is regressed on either a single intercept (in the case of the single mean model), or a switching intercept (for the switching means model). The optimal switch date for the latter model is estimated using maximum likelihood. In the first two columns individual abnormal returns are used as dependent variables. In the last three columns individual securities are first grouped into portfolios, which are then treated as single securities. Two different criteria are used in forming these portfolios: *day-of-listing*, which groups together securities with identical listing dates, and *window-overlap*, which groups together securities with overlapping 11-day event windows. The number of observations corresponding to each grouping method (N) is shown at the top of each column. t-statistics are shown in brackets under each coefficient. The last row of the table shows the significance of the switching means model compared to the single mean model (λ), which is computed as twice the difference between the log likelihood of the switching means model and that of the single mean model.

			Portfolio Grouping Method		
	Individual Listings		Day of Listing		Window Overlap
	Event Window = 11 days (N = 1924)	Event Window = 3 days (N = 1927)	Event Window = 11 days (N = 877)	Event Window = 3 days (N = 877)	Event Window = 11 days (N = 214)
Single mean model					
Mean CAR for the entire sample	−0.971%	−0.101%	−1.119%	−0.141%	−0.793%
	[−4.44]***	[−0.86]	[−3.60]***	[−0.82]	[−2.25]**
Log likelihood	1778.83	2987.28	847.76	1370.93	331.10
Switching means model					
Optimal switch date	80/12/12–81/06/29	81/07/14–81/07/15	81/07/13–81/07/14	81/07/13–81/07/14[a]	81/07/15–81/08/28
95 percent confidence interval around optimal switch date	80/08/08–82/01/26	80/07/25–82/03/10	80/07/23–82/10/22	80/08/01–82/12/22[b]	80/01/28–82/06/11
Pre-switch mean CAR	2.470%	0.851%	2.547%	1.023%	2.582%
	[5.31]***	[2.70]***	[2.68]***	[1.95]*	[3.37]***
Post-switch mean CAR	−1.509%	−0.252%	−1.548%	−0.279%	−1.593%
	[−6.29]***	[−2.00]**	[−4.75]***	[−1.54]Δ	[−4.27]***
Log likelihood	1798.34	2992.55	855.98	1373.67[c]	342.59
λ	39.0***	10.54***	16.45***	5.48**	22.97***

***, **, * Indicate statistical difference from zero at the 1, 5, and 10 percent levels, respectively.

[a]This switch date represents a local maximum of the likelihood function, which is *not* statistically different from the global maximum occurring in June 1995 ($\lambda = 1.94$, $p \approx 0.2$).

[b]This represents an 80 percent confidence interval around the local maximum. Using a 95 percent confidence interval would result in a much wider spread, making it more difficult to interpret the results.

[c]This value corresponds to the local maximum.

and five-day windows. The three-day results, shown in the table's fourth column, present a slight departure from the established pattern. In that case the global maximum of equation (1)'s log likelihood function occurs in June 1995 instead of the usual 1981 period. A closer examination of this likelihood function, however, reveals the presence of a second, local maximum in July 1981, which is not statistically different from the global maximum ($\lambda = 1.94, p \approx 0.2$). For this reason, and in order to preserve comparison with the benchmark case, I present results in this column based on the local (rather than global) maximum. Once again, the familiar pattern emerges, except that the post-switch CARs are only marginally significant. Still, the pre- and post-switch mean CARs differ from one another at the 5 percent confidence level ($t = -2.34$). The last column in Table IV shows that results obtained with 11-day CARs computed using the window-overlap grouping method closely mirror those obtained in the benchmark case. As I decrease the length of the event window toward three days, I find no departures from the result pattern already established.

The overall evidence in Table IV overwhelmingly supports H1, the hypothesis that the price effect of option introductions switches from positive to negative in the early part of the 1980s. Although the exact switch date varies with the methodology employed, the point estimates are clustered in a narrow range between late-1980 and mid-1981, and the boundaries of the 95 percent confidence intervals extend from early-1980 to late-1982. While the pre-switch results are consistent with previous studies, the post-switch negative CARs do contrast with the insignificant CARs obtained by Detemple and Jorion (1990) during the April 1982 to December 1986 period (CAR = 0.28%, $t = 0.32$). For this exact same period, I find 11-day CARs equal to −0.79 percent ($t = -1.28$).[9]

D. Switching Means versus Linear Time Trend: A Comparison of Two Plausible Models

The negative time pattern emerging from Table II suggests a second plausible alternative to the null hypothesis: a gradual temporal decrease in the price effect of options, rather that an abrupt regime switch. Formally:

[H2]: *The price effect of options, though positive in the early years of options trading, becomes gradually negative at a constant rate.* That is, the price effect follows a statistical distribution having a time-variant mean, which decreases through time in a linear fashion.

To test the truth of H2 against H0 I estimate the following linear time trend model:

$$\mathrm{CAR}_i = \alpha_0 + \alpha_1 \mathrm{TIME} + \varepsilon_i, \tag{3}$$

[9] One possible reason for this discrepancy could be the fact that I use 191 observations to form this estimate, compared to the 99 observations used by Detemple and Jorion (1990).

where TIME is the number of calendar days that, at the time of option introduction, have elapsed since the date of the first option contract initiation on the CBOE (April 26, 1973). A significantly negative α_1 coefficient indicates rejection of H0 in favor of H2. Testing H2 against H1, however, cannot be accomplished with likelihood ratio techniques, because equation (3) is not nested within specification (1) (or vice versa). I therefore use a J test developed by Davidson and MacKinnon (1981). Given any two general specifications $Y = \Phi(x)$ and $Y = \Gamma(z)$, the test consists of estimating the following hybrid supermodel:

$$Y = \gamma\Phi(x) + (1-\gamma)\Gamma(z) + \varepsilon_i. \tag{4}$$

Theoretically, a test of $\gamma = 1$ provides evidence in support of the former specification and $\gamma = 0$ supports the latter. However, since γ cannot be directly estimated from equation (4), Davidson and MacKinnon (1981) propose a two-step procedure in which each specification is tested against the other. The first step consists of estimating $Y = \Phi(x)$ and using its fitted values $F(x)$ in equation (4):

$$Y = \gamma_1 F(x) + (1-\gamma_1)\Gamma(z) + \varepsilon_i. \tag{5}$$

In the second step the procedure is repeated for the alternative model $Y = \Gamma(z)$:

$$Y = (1-\gamma_2)\Phi(x) + \gamma_2 G(z) + \varepsilon_i, \tag{6}$$

where $G(z)$ are the fitted values of $\Gamma(z)$. Evidence in support of $Y = \Phi(x)$ is obtained when γ_1 is significantly positive, but statistically indistinguishable from one, *and* when γ_2 is statistically indistinguishable from zero. By contrast, evidence in support of $Y = \Gamma(z)$ is obtained when $\gamma_1 \approx 0$ and $\gamma_2 \approx 1$ in a statistical sense. To understand the intuition behind this test, suppose $Y = \Phi(x)$ is the true model, so that $Y = \Gamma(z)$ is necessarily false. In this case, the fitted values $G(z)$ from the false model should have no explanatory power for Y, in the presence of the true model. Statistically, this is verified when γ_2 is not significantly different from zero in equation (6). Additionally, if $Y = \Phi(x)$ is true, its own fitted values $F(x)$ should be the only factor accounting for variations in Y, in the presence of the false model. This condition is verified by testing the hypothesis that $\gamma_1 = 1$ in equation (5), which requires that γ_1 be significantly positive and that $(1 - \gamma_1)$ be statistically insignificant.

A major weakness of the J test is that it may lead to "inconclusive" results, such as rejecting both models, or rejecting neither. This problem, however, is confined to finite samples. Davidson and MacKinnon (1981) show that if $Y = \Phi(x)$ is the true model, the probability that γ_1 differs significantly from zero approaches one as the number of observations approaches infinity. Another shortcoming is that γ_1 and γ_2 are only asymptotically normally distributed. The J test is therefore appropriate only when used with large samples.

I use this test to discriminate between H1 and H2, with equation (1) in place of $Y = \Phi(x)$, and equation (3) in place of $Y = \Gamma(z)$. Rejecting the linear time trend model (H2) in favor of the switching means hypothesis (H1) requires that γ_1 be statistically close to one (and different from zero), and that γ_2 be close to zero. Table V presents the results. In the upper half of the table I present estimates for the linear time trend model described by equation (3). In the lower half I present estimates of the γ_1 and γ_2 coefficients. The horizontal axis is once again divided into five columns, corresponding to various combinations of grouping methods and event windows. The benchmark results are presented in the first column. Although the coefficient of TIME in equation (3) is significantly negative, the J test shows that the linear time trend model H2 can be convincingly rejected in favor of the switching means model H1, because γ_1 is close to one (and statistically different from zero), while γ_2 is small and statistically insignificant. This result holds for all possible combinations of event windows and grouping methods, the most representative of which are illustrated in the remaining columns of Table V.[10]

II. Possible Causes for the Switch in the Price Effect of Options

The evidence thus far presented strongly supports the regime switch hypothesis H1, that pre-1980 and post-1980 option introductions had significantly different mean effects on the underlying stock's price. Perhaps this change reflects some observable characteristics of the optioned firms, rather than a change in the underlying economics of option introduction.

Two such characteristics are the AGE and SIZE of underlying stocks, which appear to follow a negative time trend throughout the sample period (Table I). In recent years, most optioned stocks are small and relatively less seasoned. For this particular group, the costs of establishing short positions may be prohibitive before options are listed, to the extent that bearish investors who do not own the stock are unable to borrow it. When options are introduced, these short sale restrictions are effectively eliminated. By contrast, the more established equities of the pre-1980 period may not be subject to any short sale restrictions before options, since borrowing the stock may be relatively easy.

Diamond and Verrecchia (1987) show that when option listings eliminate binding short sale constraints, underlying stock prices adjust more quickly to adverse information. Before options, short-sale-restricted stocks are slower

[10] Although the J test could be inconclusive in finite samples, I do not encounter such situations in this study, owing perhaps to the relatively large sample size employed.

Table V

A Statistical Comparison of the Switching Means and Linear Time Trend Models

The dependent variable is the cumulative abnormal return (CAR), computed using the Brown and Warner (1985) methodology, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. The dependent variable is regressed on a linear time trend variable (TIME). TIME is computed as the number of calendar days that, at the time of option introduction, have elapsed since the date of the first option contract initiation on the CBOE (April 26, 1973). The estimation method is OLS, with standard errors computed using a White heteroskedasticity-consistent variance-covariance matrix. In the first two columns of the table, individual abnormal returns are used as dependent variables. In the last three columns individual securities are first grouped into portfolios, which are then treated as single securities. Two different criteria are used in forming these portfolios: *day-of-listing*, which groups together securities with identical listing dates, and *window-overlap*, which groups together securities with overlapping 11-day event windows. The number of observations corresponding to each grouping method (N) is shown at the top of each column. t-statistics are shown in brackets under each coefficient. The significance of the switching means model compared to the linear time trend model (γ_1) is established using the Davidson and MacKinnon (1981) methodology. A theoretical value of $\gamma_1 = 1$ indicates that the data are best described by the switching means model. Note that all estimated γ_1 coefficients in this table are statistically different from zero, and none is statistically different from one. The significance of the linear time trend model compared to the switching means model (γ_2) is established using the same methodology. None of the γ_2 coefficients are statistically different from zero, indicating that the linear time trend model is outperformed by the switching means model.

	Individual Listings		Portfolio Grouping Method: Day of Listing		Portfolio Grouping Method: Window Overlap
	Event Window = 11 days (N = 1924)	Event Window = 3 days (N = 1927)	Event Window = 11 days (N = 877)	Event Window = 3 days (N = 877)	Event Window = 11 days (N = 214)
Estimates of the linear time trend model					
Intercept	0.0224	0.0063	0.0210	0.0070	0.0225
	[4.52***]	[2.51***]	[2.78***]	[1.88*]	[2.71]***
TIME ($\times 10^{-5}$)	−0.55	−0.13	−0.53	−0.14	−0.61
	[−6.26]***	[−2.72]***	[−4.03]***	[−2.00]**	[−4.01]***
γ_1	0.7609	1.1484	0.8408	1.1664	0.9067
	[3.35]***	[2.54]***	[2.62]***	[2.24]**	[2.73]***
γ_2	0.3229	−0.2436	0.2347	−0.3009	0.1421
	[1.18]	[−0.36]	[0.57]	[−0.36]	[0.36]

***, **, * Indicate statistically different from zero at the 1, 5, and 10 percent levels, respectively.

to adjust to bad news than they are to good news. When options are listed, the speed of adjustment to bad news increases, lowering the average price of the affected underlying securities. To better understand this mechanism, assume that for a cross section of short-sale-constrained securities, those carrying good news are trading at their true value while most of those carrying bad news are temporarily trading above their true value. If short sale constraints are never released, the overpriced securities will *slowly* adjust toward their true value as more and more informed traders who actually own the stock begin to sell. However, when option listings cause short sale constraints to be *simultaneously* released for the cross section of securities, the market value of the overpriced securities *quickly* drops toward the true value. Because stocks carrying good news are already correctly priced, the net effect of options introduction is a short-term decrease in the average price of the cross section of underlying securities.

This short sale hypothesis may provide a plausible explanation for the negative CAR time pattern documented in the previous section: Though options may decrease the price of *short-sale-constrained* securities, they leave unaffected the price of underlying stocks that *are not* short-sale restricted. Because the severity of short sale restrictions is likely to be greater for stocks listed after 1980 (due to the small AGE and SIZE of underlying firms), the price effect of option introductions is more likely to be negative during that period.[11]

AGE and SIZE are not the only observable characteristics that changed throughout the sample period. Further examination of Table I also reveals temporal changes in the type of option contracts listed, as well as in the trading place for options and their underlying stocks. Since pre-switch options are predominantly call-only contracts, it is possible that the switch reflects the advent of joint put/call contracts in the post-1980 period. It is also conceivable that options traded on the AMEX, Philadelphia, and Pacific exchanges somehow affect the underlying stock differently from options traded on the CBOE, as may options that are simultaneously listed on more than one exchange. Additionally, stocks trading on the Nasdaq exchange (which only became optioned in 1982) may respond differently than stocks trading on the NYSE or AMEX.

To evaluate these potential effects, I estimate the following augmented versions of equations (1), (2), and (3):

$$\mathrm{CAR}_i = I_i\alpha_1 + (1 - I_i)\alpha_2 + \beta_1 X + \varepsilon_i \tag{1'}$$

$$\mathrm{CAR}_i = \alpha_0 + \beta_2 X + \varepsilon_i \tag{2'}$$

$$\mathrm{CAR}_i = \alpha_0 + \alpha_1 \mathrm{TIME} + \beta_3 X + \varepsilon_i . \tag{3'}$$

[11] I thank Bartley Danielsen for suggesting this short-sales hypothesis.

In each case the set of independent variables X includes various combinations of the underlying firm's AGE and SIZE, as well as dummy variables for various option exchanges, call-only contracts, and Nasdaq-traded stocks.

If the switch is fully explained by the observable characteristics included in the X matrix, equation (2′) will provide the best fit for the data, and at least one of its β_2 coefficients will be statistically significant. Further, the switching intercept in equation (1′) will be insignificant. By contrast, if none of the X variables account for the switch, equation (1′) will provide the best fit for the data. Also, the only significant parameters within that equation should be the two switching intercepts, which should remain associated with the 1981 period. The β_1 coefficients should be statistically insignificant. As in the previous section, I use the likelihood ratio test to compare specifications (1′) and (2′), and the J test to compare specifications (1′) and (3′).

Table VI presents the benchmark results. In the top section I show the date and the estimates of the switching intercept in equation (1′), as well as the values of the corresponding β_1 coefficients. In the bottom section I provide a test for multicollinearity, the regression's adjusted R^2, and the familiar λ statistic and γ coefficients. Along the horizontal axis, I explore various combinations of independent variables. The first column of Table VI shows that the switch cannot be related to the introduction of joint put/call contracts in the post-1981 period because equation (1′) continues to outperform both (2′) and (3′) in the presence of a dummy for call-only listings. Moreover, the optimal switch date does not change, and the coefficient of the Call-only dummy is statistically insignificant. Similarly, the second column in Table VI shows that the switch is unrelated to the trading place for options or their underlying stocks.[12] I also investigate and dismiss the hypothesis that the switch is related to the advent of options listed on multiple exchanges in the post-1980 period (not shown).

I next examine the hypothesis that the switch in the CAR might somehow reflect the fact that recently optioned stocks are more difficult to short without options. If the stock's AGE and SIZE are reasonable proxies for the difficulty of establishing short positions, I expect that equation (2′) will provide

[12] This conclusion is strengthened by additional tests in which equation (2) is rejected in favor of equation (1) for each option exchange subsample, as well as for the subsample of stocks trading only on the NYSE and AMEX. In all instances the switch date remains essentially unchanged. Two somewhat puzzling observations in this column are the significant coefficients of Nasdaq-listed stocks and NYSE-listed options dummies. Further analysis, however, shows that the sign of the Nasdaq coefficient is not robust to changes in the length of the event window. More puzzling is the positive coefficient on NYSE-listed options, which remains generally robust to changes in the length of the event window. Since NYSE options were introduced exclusively in the post-switch period, I expect their mean CARs to be significantly negative. Yet the mean CAR for this subsample is barely negative (−0.10 percent) and statistically insignificant ($t = -0.15$). It appears, however, that this result is driven by one outlier where the underlying firm experiences a positive 50.6 percent event window abnormal return. This is corroborated by observing that the *median* CAR for the NYSE options subsample is of the same order of magnitude as in the benchmark case (−1.06 percent), although the significance level is very marginal at best (sign rank statistic = −9, $p = 0.22$).

Table VI

Switching Intercept–Linear Cross-Sectional Model of the 11-day CAR

The dependent variable is the cumulative abnormal return (CAR) using individual listings (no portfolio groupings). CARs are computed using the Brown and Warner (1985) methodology, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. Independent variables are: Log(SIZE), the natural logarithm of the firm's market capitalization, expressed in terms of 1983 dollars; Log(AGE), the natural logarithm of the number of calendar days that have elapsed since the firm first became publicly traded; Call-only dummy, a dummy variable equal to one when only call contracts are being listed, and equal to zero when both calls and puts are listed simultaneously; NASD stock dummy, a dummy variable equal to one when the underlying stock trades on the Nasdaq exchange at the time of option listing, and equal to zero otherwise; and AMEX, PHLX, PSEX, NYSE dummies, variables equal to one when the option trades on the American, Philadelphia, Pacific, and New York option exchanges respectively, and equal to zero otherwise (when all four dummies equal zero, the option trades on the Chicago Board of Options Exchange). The model is linear in these independent variables, but allows for a switching intercept. The optimal switch date of the intercept is estimated using the maximum likelihood method. Values shown in the table correspond to global maxima of the log likelihood function. The independent variables are estimated using the OLS method, with standard errors computed using a White heteroskedasticity-consistent variance-covariance matrix. Across the horizontal axis, the table explores various combinations of independent variables and sample selections. The number of observations corresponding to each model specification is shown at the top of each column. *t*-statistics are shown in brackets below each entry. The table also reports a test for collinearity, which is the Belsley condition number described in Greene (1993). This number is computed as the square root of the ratio of the largest to the smallest characteristic root of the $X'X$ matrix. Generally, a condition index in excess of 20 indicates collinearity problems. (Greene (1993), 269.) The significance of the switching intercept model compared to an otherwise identical single intercept model (λ) is computed as twice the difference between the log likelihood of the two models. The switching intercept model is also compared with a linear time trend model, in which the switching intercept is replaced with a single intercept and a time trend variable (TIME). TIME is defined as the number of calendar days that, at the time of option introduction, have elapsed since the date of the first option contract initiation on CBOE (April 26, 1973). The other independent variables remain unchanged. The significance of the switching intercept model compared to the linear time trend model (γ_1) is established using the Davidson and MacKinnon (1981) methodology. A theoretical value of $\gamma_1 = 1$ indicates that the data are best described by the switching intercept model. Note that all estimated γ_1 coefficients in this table are statistically different from zero, and none are statistically different from one. The significance of the linear time trend model compared to the switching intercept model (γ_2) is established using the same methodology. None of the γ_2 coefficients are statistically different from zero, indicating that the linear time trend model is outperformed by the switching intercept model.

	Model No.					
	(1)	(2)	(3)	(4)	(5)	(6)
Sample	Full Sample				Older Firms Only (AGE > 5 years)	Larger Firms Only (SIZE > median)
No. of observations	1924	1924	1922	1922	1183	958
Optimal switch date	80/12/12–81/06/29	80/12/12–81/06/29	80/12/12–81/06/29	80/12/12–81/06/29	80/12/12–81/06/29	80/12/12–81/06/29
Pre-switch intercept	0.0193	0.0242	0.00820	0.0662	0.0248	0.0233
	[2.89]***	[4.63]***	[0.14]	[0.99]	[5.12]***	[5.05]***
Post-switch intercept	−0.0164	−0.0133	−0.0296	0.0292	−0.0120	−0.0100
	[−5.73]***	[−3.12]***	[−0.55]	[0.45]	[−4.31]***	[−3.39]***
Log(SIZE)	—	—	0.00022	−0.00230	—	—
			[0.08]	[−0.73]		
Log(AGE)	—	—	0.00133	0.00031	—	—
			[0.83]	[0.19]		
Call-only dummy	0.0055	—	—	0.00415	—	—
	[1.12]			[0.72]		
NASD stock dummy	—	−0.0100	—	−0.0107	—	—
		[−1.97]**		[−1.77]*		
AMEX option dummy	—	−0.00274	—	−0.00261	—	—
		[−0.50]		[−0.48]		
PHLX option dummy	—	0.00854	—	0.00886	—	—
		[1.58]		[1.63]		
PSEX option dummy	—	−0.00044	—	−0.00049	—	—
		[−0.08]		[−0.09]		
NYSE option dummy	—	0.0152	—	0.0158	—	—
		[2.07]**		[2.13]**		
Test for collinearity (condition number)	2.52	3.64	48.9##	68.6##	N/A	N/A
Adjusted R^2	0.0196	0.0238	0.0190	0.0229	0.0302	0.0326
λ	22.1***	30.5***	27.6***	22.1***	37.2***	32.7***
γ_1	0.8811	0.8960	0.8161	0.9305	0.7053	0.8887
	[3.20]***	[3.68]***	[3.39]***	[3.47]***	[2.65]***	[2.82]***
γ_2	0.1882	0.1587	0.2829	0.1185	0.3770	0.1488
	[0.53]	[0.49]	[0.88]	[0.33]	[1.22]	[0.40]

***, **, * Indicate statistically different from zero at the 1, 5, and 10 percent levels, respectively
Indicates a possible multicollinearity problem present in this specification.

the best fit for the data when these two variables are included in the X matrix. The results are presented in the third and fourth columns of Table VI. The switching intercept model (1′) still provides the best fit for the data, but the coefficients of the switching intercept are now insignificant. As shown in the table's bottom portion, this problem is likely due to the presence of multicollinearity between AGE, SIZE, and the switching intercept. Without additional testing, it is not possible to determine whether the switch continues to be significant after accounting for the underlying stock's AGE and SIZE. I use two tests to overcome this multicollinearity problem.[13]

The first consists of reestimating equations (1) through (3) for the subsample of older firms, and also for the subsample of larger firms. A firm is deemed "old" if it has been publicly traded for at least five years at the time of its option introduction, and "large" if its SIZE is higher than the median. If the switch is unrelated to the advent of younger, smaller firms in the latter part of my study period, the optimal switch date in equation (1) should remain in the vicinity of 1981, and equation (1) should continue to outperform equations (2) and (3) for each of the two subsamples. The results presented in the fifth and sixth columns of Table VI readily support this conjecture. Even in the case of more seasoned and, alternatively, larger firms, the price effect switches from significantly positive to significantly negative around the usual 1981 period. Equation (1′) also continues to provide the best fit for the data when compared to the two alternative specifications.

For the second test I compare the CARs of each sample firm in the pre-switch period with the CARs of one control firm in the post-switch period that is closest in AGE and SIZE. If the regime switch is indeed unrelated to the AGE and SIZE of the underlying firm, CARs should be positive in the pre-switch sample, and negative in the post-switch sample, and the mean pairwise difference between the post- and pre-switch CARs should be significantly negative. Table VII presents these results. The first column in Table VII presents the benchmark results for 11-day CARs computed from individual listings. Even after matching by AGE and SIZE, CARs continue to switch from significantly positive to significantly negative in 1981. Additionally, the mean pairwise difference between the two subsamples is negative at better that the 1 percent level. As I decrease the width of the event window toward three days, I find results that are qualitatively similar to the benchmark case, except for the three-day event window when post-switch CARs are negative although insignificant. Results for the three-day window are shown in the second column of Table VII. Even in this case, however, the mean CAR difference between post- and pre-switch samples remains significantly negative.

The evidence presented in this section shows that the switch in the price effect of options is not related to the type or trading place of the typical option contract, nor to the AGE, SIZE, or trading place of the typical under-

[13] I thank the referee for suggesting these two approaches.

Table VII

The Effect of AGE and SIZE on the CAR Switch: A Control Firm Approach

This table compares the cumulative abnormal returns (CARs) of each sample firm in the pre-switch period with the CARs of a control firm in the post-switch period that is closest in AGE and SIZE. AGE is the number of calendar days that, at the time of option introduction, have elapsed since the firm first became publicly traded. SIZE is the value of the firm's market capitalization, expressed in constant (1983) dollars. To match each sample firm with a suitable control firm I first divide all firms into "old" and "young" AGE groups, using five years as a cutoff AGE. For each sample firm in the pre-switch period I then select a control firm from the post-switch period belonging to the same AGE group, and whose SIZE is closest to that of the sample firm without being different by more than 20 percent. From the 258 pre-switch observations, only 213 meet this 20 percent criterion (N). The remaining 45 (mostly firms listed in 1973–1974) were too large to find a suitable "match" in the post-switch period. CARs are computed using the Brown and Warner (1985) methodology, with a benchmark model estimated over the [L − 100 to L − 6] period (L is the day of option listing). Market returns used are the combined, value-weighted NYSE–AMEX–Nasdaq index. The appropriate test statistics are shown in brackets below the point estimate, along with significance levels. For mean returns the test statistic shown is the usual t-statistic. For median returns I show the Wilcoxon sign rank statistic.

	CAR Estimation Window	
	11 days	3 days
Pre-switch period: 1973–1980 (N = 213)		
Mean firm SIZE ($million 1983)	1,553	1,553
Median firm SIZE ($million 1983)	1,048	1,048
Mean underlying firm AGE (years)	26.3	26.3
Median underlying firm AGE (years)	19.6	19.6
Mean CAR	0.0298	0.0099
	[5.57]***	[3.92]***
Median CAR	0.0203	0.0089
	[25.5]***	[19.5]***
Post-switch period: 1981–1995 (N = 213)		
Mean firm SIZE ($million 1983)	1,590	1,590
Median firm SIZE ($million 1983)	1,052	1,052
Mean underlying firm AGE (years)	23.2	23.2
Median underlying firm AGE (years)	19.2	19.2
Mean CAR	−0.0093	−0.0033
	[−1.88]*	[−1.34]
Median CAR	−0.0094	−0.0026
	[−18.5]**	[−11.5]Δ
Pairwise comparison of the two periods (N = 213)		
Mean (post-switch CAR less pre-switch CAR)	−0.0391	−0.0132
	[−5.27]***	[−3.59]***
Median (post-switch CAR less pre-switch CAR)	−0.0322	−0.0081
	[−32.5]***	[−14.5]*

***, **, *, Δ Indicate statistically different from zero at the 1, 5, 10, and 15 percent levels, respectively.

lying stock. It should be noted, however, that these independent variables are only several of the possible firm characteristics that could be responsible for the difference in the price effect of options. To assess the robustness of this conclusion I repeat much of this section's analysis with nine-, seven-, five-, and three-day event windows. I also group observations into day-of-listing portfolios, by averaging CARs, SIZE, AGE, and all dummy variables across all firms with identical option introduction dates. I then repeat most of my tests with CARs computed using decile-based size-matched portfolios instead of the value-weighted market index. Overall, the results are remarkably robust to any such changes in methodology: Regardless of the event window, grouping method, and CAR computation method, specification (1′) continues to provide the best fit for the data, and the optimal switch date remains in the neighborhood of June 1981. Corrections for multicollinearity also yield the same qualitative results. The only departures from this pattern occur when CARs are computed over the shorter three-day (and in one instance five-day) window, in which case the global maximum of the likelihood function in equation (1′) occurs in 1995. However, even in these instances, the likelihood function exhibits a second, *local* maximum around 1981. Using a likelihood ratio test one generally cannot reject the hypothesis that the local maximum is in fact the function's true maximum.[14] When equation (1′) is reestimated at the 1981 local maximum, the usual result pattern emerges even for the shortest event windows.

III. Discussion and Conclusion

This study demonstrates that cross-sectional variations in the price effect of option introductions are best explained by a two-regime, switching means model. The CARs are drawn from a positive-mean distribution prior to 1981, and a negative-mean distribution thereafter. Cross-sectional variations in underlying firms' AGE and SIZE are mere proxies for the time regime, and have no direct impact on CARs. Nor does the switch seem to be related to temporal patterns in the type of option contracts being listed, or to changes in the trading place for options or their underlying stocks. The switch could however reflect changes in *other* firm characteristics, which are not explored in this paper.

Although the reasons for this regime shift are not immediately apparent, one promising area for further research remains the market completion hypothesis of Detemple and Jorion (1990), who suggest that the introduction of stock index options on April 21, 1982 effectively completed the markets and

[14] For example, when I use three-day CARs based on the market model and individual listings, I find that the global maximum of the likelihood function of equation (1′) moves to October 1995 when AGE and SIZE are included in the X variables. Nonetheless, a local maximum is present around the familiar 1981 period and the hypothesis that the "true" maximum is in fact the 1981 maximum cannot be rejected at conventional significance levels ($\lambda = 1.80$, $p \approx 0.25$).

allowed investors to span the relevant state space. Under this hypothesis the price effect of options is positive prior to the index option introduction, but should subsequently vanish. Consistent with this hypothesis, the authors find positive CARs prior to April 1982, and statistically insignificant CARs after that date, although the period they are able to analyze ends in 1986.

Another possible reason for the switch stems from an examination of the regulatory environment of standardized options trading. After an SEC-imposed moratorium on new option listings during 1977 to 1979, a study recommended a number of substantial changes in the operating procedures of option exchanges. These changes were implemented during 1980 and 1981. They generally imposed more stringent margin restrictions for option dealers, and severely limited the use of deceptive sale practices in marketing options to unsophisticated investors. Under this hypothesis, the positive effect during 1973 to 1980 might reflect some sort of regulatory inefficiency during that period. However, the price effect should vanish with the implementation of new regulations in 1980–1981.

Under both the market completion and regulatory change hypotheses, the price effect of options should be close to zero after 1981 or 1982. Although this implication is not supported by the results obtained in this paper, one cannot necessarily conclude that both hypotheses should be rejected. Rather, this fact merely suggests that if either of these two hypotheses is true, a different factor must be responsible for the negative CARs observed in the post-1980 period.

One such factor remains the dissemination of unfavorable information by informed traders who found it difficult to establish short positions before options were introduced. Under this explanation, a negative information effect would affect stock prices adversely *throughout* the sample period when option trading is introduced, but the measured positive reaction before 1981–1982 reflects a dominance of the market completion or regulatory inefficiency effects during that period. The lack of a cross-sectional relation between CARs and the underlying firms' AGE and SIZE is not necessarily inconsistent with this information hypothesis, because AGE and SIZE might not adequately reflect the extent to which shorting is difficult in the absence of options. Future research could try to identify better proxies for the severity of short sale restrictions, in an effort to increase our understanding of the relation between short sale constraints and security prices.

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