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Do Newly Listed Derivatives Affect the Market Risk Premium in a Thin Stock Market?

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Abstract. This study examines the effects on the stock market unitary risk premium and volatility associated with the listing of stock and stock index derivatives in Switzerland. Based on a univariate GARCH (1,1) specification of the stock index variance and a time-varying unitary risk premium representation, we can reject the hypothesis that stock and stock index derivatives listings do not affect the total risk premium. Contrarily to previous empirical evidence, we find that derivatives listings affect both the conditional market returns' variance and the unitary risk premium through structural shocks. The gradual market completion hypothesis is further corroborated in that, cumulatively, the three stock and stock index options futures derivatives listings reduced the unitary risk premium while the marginal impact of each successive listing decayed.

Key words: futures, non-redundancy, options, time-varying risk premium, volatility.

JEL Classification: G12, G14.

1. Introduction

It is well known that most contingent claims pricing models are built under the assumption that the derivative is a redundant asset. Therefore, its listing and subsequent trading should not affect the return or volume distributions of the underlying asset. This study tests another dimension of the redundancy property of the listing of stock and stock index options and futures on the Swiss stock market namely whether they can lead to shifts in the representative agent's preferences. More precisely, its objective is to assess whether the listing of standardized derivatives affects the unitary stock market premia. Under the null hypothesis, it is well

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known that derivatives listing and subsequent trading should not alter the return distribution of the underlying stock or stock index. Most theoretical and empirical papers have challenged the latter statement under the premise that derivatives complete the markets and thus affect the underlying asset's price distribution only. For instance, several empirical papers (see Conrad (1989), Detemple and Jorion (1990)) have documented that the introduction of traded call options leads to significant positive abnormal excess returns while the converse significant negative excess returns follow the introduction of listed put options (see Damodaran and Lim (1991)). Secondly, the impact of option and futures listing on the volatility of the underlying stock or stock index was examined. Studies were almost unanimous – irrespective of the computation method, the time period or the data sample chosen – in acknowledging a decline in the total variance or systematic risk of the optioned stock. For instance, Edwards (1988a and 1988b) reports a decline in the stock market volatility after the introduction of stock index futures in 1982. On the Swiss market, Bruand and Gibson (1998) study the informational feedback effects associated with the listing of standardized stock and stock index options. Their empirical results support the gradual completion hypothesis advocated by Detemple and Jorion (1990) in that the decrease of the variance of the blue chips stocks and of the Swiss Market Index (SMI) returns becomes less pronounced after subsequent launchings. The study by Damodaran and Subrahmanyam (1992) is unique in providing an exhaustive survey on the non-redundancy property of futures and options both from a theoretical and empirical standpoint and emphasizing elaborately the pricing and volatility effects associated with the listing and trading of derivatives in the United States.

In contrast to the abundant empirical studies previously mentioned, very few theoretical models have so far been proposed in order to explain the non-redundancy and the economic functions of financial derivatives in imperfect markets. The first early contribution on the effects of options listing in incomplete markets (see Ross (1976)) has been successfully pursued in the general equilibrium models of Detemple and Selden (1991) and Detemple and Jorion (1990). The latter authors show that if agents have different risk aversion coefficients, introducing options in the presence of several risky assets (say, on an individual stock and stock market portfolio) will increase the stock and stock indexes' prices. It will also decrease the stock volatility as a result of direct and cross market effects (as long as their returns correlation differs from zero). In addition, they show that once the first option is introduced, the listing effect for the second one is weaker given that a certain degree of market completeness has already been reached in the economy. They further report tests for options listings on the CBOE and on the AMEX which support the predictions of their theoretical model.

In a market with frictions (such as informational asymmetries, monetary transaction costs, short selling restrictions or leverage constraints), the economic role of derivatives extends beyond pure risk sharing. Indeed, market participants will also value their asset allocation and price discovery functions and, as a result, price them as non-redundant claims. For instance, investors may choose to reveal their private information primarily through derivatives rather than cash markets transactions, thus increasing the precision and speed of information gathering after options or futures are listed. Along these lines of thoughts, Grossman (1988) shows that, in the presence of asymmetric information, "real" and "dynamically replicated" derivative securities are imperfect substitutes since the former options (say, a European put) prices convey additional and more precise information to market participants (than a dynamic portfolio insurance strategy). Along the same vein, if derivative securities are imperfectly correlated with their dynamic cash market replicating strategies, their trading whether for asset allocation, arbitrage or informational purposes will affect the underlying asset's return characteristics. Stein (1989) develops a model in which the destabilizing – due to speculative trading by less informed investors – effect of introducing options dominates their stabilizing risk sharing function and results in an increase in the market's volatility. A related conclusion is reached by Back (1993) who shows that, in the presence of asymmetric information, stock options can no longer be priced as redundant assets since their trading conveys additional information. Relying on an extension of Kyle's (1989) model he shows that, after the listing, the volatility of the underlying stock becomes stochastic as a result of traders' heterogeneity. Indeed, traders with superior information will choose to trade out-of-the-money options. Back's result holds as long as the liquidity trades in the cash and in the derivative securities are imperfectly correlated and it applies to any derivative security with non-linear payoffs. Biais and Hillion (1994) tackle more precisely the informational efficiency of the market when non-redundant options are introduced. They foster the ambiguous consequences of option introduction since their trading conveys information whose interpretation can be less informative given the larger set of investment strategies they provide to market participants.

In all these studies, there is neither a model nor empirical evidence which examines whether the listing process actually alters equilibrium prices via a modification of the unitary market risk premia coupled – but not necessarily – with a modification of the covariance structure of asset returns. Yet, we know from several asset pricing models that the equilibrium real interest rate and asset returns prices may differ in complete and incomplete markets. For instance, Detemple and Serrat (1998) have shown that market incompleteness and liquidity constraints unambiguously lower the real interest rate. Furthermore, they show that liquidity constraints increase the Sharpe ratio (yet, not enough, to explain the equity premium puzzle).

It is the purpose of this study to empirically assess whether derivatives non-redundancy only manifests itself through structural (non-transitory) shifts in the return distribution of the underlying security and/or through a modification of agents' unitary risk premia. The model of market returns that will be used in this study is a conditional Capital Asset Pricing Model (CAPM). Indeed, on the theoretical side, this modelling is widely used in the literature and, empirically, tests seem to show that it is more justified than a consumption CAPM or a multifactor pricing model.[1] In a CAPM setting, we can express the risk premium as the relative risk aversion coefficient of the representative agent multiplied by the volatility of aggregate consumption, which could be approximated by the volatility of the stock market if the difference between the volatilities is small. We remark hence that there is a priori no reason to impose a constant unitary risk premium since agents' relative risk aversion may potentially vary with modifications in agents' endowments, in their standard consumption levels or with the business cycle. Furthermore, one could conjecture that when agents face a broader investment opportunity set through financial innovation, they will lower their precautionary savings and thus that the riskless rate will increase as a result of agents being less constrained. Following Detemple and Serrat' (1998) model, financial innovation, more specifically through the listing of new derivatives, can be perceived as a mechanism which enables agents to lower their liquidity constraints. As a result, their introduction should lower the Sharpe ratio on the market portfolio if aggregate risk aversion is a decreasing function of the consumption share of constrained individuals. Thus, the listing of options and futures may be followed by a structural decline in the unitary market risk premium. Let us however be aware that there is little theoretical guidance, which allows us to explicitly sign the effect of the launching of new options and futures contracts on the unitary risk premium. Indeed, the significance, sign, and magnitude of this effect may depend upon the type of contract being introduced (option versus futures), the type (allocational versus informational) and degree of market incompleteness and the degree of international market integration prevailing prior to their launching. In the absence of a directly applicable theoretical model, we shall thus simply test whether there is any effect subsequent to the derivatives listings on the unitary risk premium against the null hypothesis of options redundancy which assumes no modification of the unitary risk premium subsequent to any derivatives listing.

In order to test the joint effect of three subsequent stock and stock index derivatives listings, we assume that the unitary market risk premium is time-varying and will express it as a function of a set of instrumental variables. This allows us to test for a richer set of potential risk premia modifications which can occur through a shift in their constant level and/or sensitivities to the instrumental variables coefficients. We shall also investigate whether the modification of the unitary risk premium strengthens or dominates the variance reduction effects as reported in the previously cited studies and finally test for gradual market completion subsequent to consecutive derivatives listings. The empirical study will be conducted on the Swiss market in order to examine this specific issue for a thinly traded and concentrated stock market. Hence, assuming a time-varying unitary risk premium specification coupled with a univariate GARCH (1,1)-in-mean model for the Swiss Bank Corporation General (SBC) Index excess returns, we analyze the individual and joint effects of the three stock and stock index derivatives listings which took

[1] See Campbell, Lo and MacKinlay (1997) for a comprehensive overview of each pricing model and its associated empirical results.

place on the Swiss stock market between the 19th May 1988 and the 9th November 1990. We further rely on previous empirical results by Bruand and Gibson (1998) that acknowledged a significant reduction in the unconditional variance of the SBC General Index's (and of the Swiss Market Index, SMI) returns after the three derivatives listings and used a GARCH (1,1) specification to show that cumulatively, the three listings also significantly reduced the constant of the SMI (Swiss Market Index's) returns' conditional variance process.

In this study, we show that the three derivatives listings also affect the unitary market risk premium. Indeed, they cumulatively reduce the unitary risk premium by 5.5%.The feedback effects documented in previous studies that only focused on stock and stock index variance effects due to market completion through derivatives thus only offer a partial representation of the total listings' effects. It appears that the derivatives listings jointly affect the market unitary risk premium and the GARCH (1,1) variance specification. One should however treat these empirical results with caution and stress the joint hypothesis nature of the test conducted in this study. That is, we assume the validity of the asserted parameterization while investigating the redundancy of derivatives launchings. In a related context, Scruggs (1998) studies the misspecification bias induced by the estimation of an univariate CAPM model. He shows that if the true model is the intertemporal one state variable CAPM, as developed by Merton (1973), then estimation of the "partial" risk-return relation in the univariate model can be misleading.[2]

We also tackle the question of how many instruments and which ones should be included in the information set to correctly represent the unobserved time-varying unitary risk premium. It seems that the number of necessary instruments is mainly driven by both the volatilities and the cross-correlations in the instrumental variables series. Since we do not expect the market risk aversion parameter variations to be extreme but rather smooth, it is possible that adding instrumental variables may perhaps not improve the informational content of the set of instruments but may lead to a less volatile weighted average of the instruments and thereby improve the likelihood of the parameterization. The latter hypothesis is also tested and leads to the following conclusion: given a smoothing method,[3] there exists an "optimal" smoothing parameter which meaningfully improves the likelihood when applied to the instrumental variables series. One should however notice that the economic conclusions are somewhat modified compared to the ones drawn on raw instruments. It is moreover difficult to determine if these modifications are mainly due to the smoothing method used (in this study, a Hodrick-Prescott filter) or due to the economic content of the smoothed instrumental variables series.

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[2] The sign of the bias in the estimated unitary risk premium varies from significatively negative to significatively positive depending on, firstly, the sign of the partial derivative of the indirect utility function with respect to the wealth and the state variable and, secondly, the sign of the covariance between the market variance and the market covariance with the state variable $\left(\text{cov}\left(\sigma_{M,t}^2, \sigma_{MF,t}\right)\right)$.

[3] The issue of the optimal smooting method is left for future research.

The structure of the paper is the following: in Section 2, the characteristics of the Swiss stock and derivatives markets are presented. In Section 3, we present the data and the specification of the conditional asset pricing model which rests on a GARCH (1,1) specification for the variance of the stock market returns. In Section 4, we examine whether the listing of individual stock options, stock index options and futures did individually and jointly affect the unitary time-varying market risk premium. We further examine the role of the unitary premium smoothening on the latter conjecture. Section 5 concludes the paper and discusses some avenues for further research on the interaction between derivatives and cash markets.

2. General Description of the Swiss Stock and Derivatives Markets

The Swiss stock market is highly concentrated and its trading activity is segmented by firm size as suggested by the fact that the 10 largest firms account for approximately 80% of its total capitalization. Swiss corporations can issue bearer, registered and non-voting stocks, which all display different voting privileges. In particular, registered shares have superior voting privileges since they can be issued at a lower par value than bearer shares. They still obey to the "one share – one vote" rule. Until the 1992 revision of the corporate law, Swiss companies could issue restricted registered shares whose ownership transfer could not be enforced without the company's formal agreement. This mechanism was used in order to monitor and restrict foreign ownership and control. For these reasons, bearer shares have generally been considered as more liquid, especially prior to November 1988[4] where they traded at a premium ranging between 20 to 40% over registered shares issued by the same firms.

The SOFFEX (Swiss Options and Financial Futures Exchange) is a fully computerized continuous trading derivatives' exchange which is supported by less than twenty official market makers. It started trading American options on 11 individual stocks on May 19th, 1988. The launching of a standardized derivatives market was intended to boost liquidity in the underlying cash market. The introduction of individual stock options was followed by the launching of SMI (Swiss Market Index[5]) American options and futures contracts respectively on December 7th, 1988 and November 9th, 1990. It is interesting to observe that, contrarily to the US experience, stock options were launched first. Initially, trading activity was thin and irrational early exercise policies were acknowledged. Thus, SOFFEX decided to convert the SMI index options to European options during the second half of 1990.

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[4] On November 18th, 1988, Nestlé opened its equity ownership to foreign investors allowing them to buy its Restricted Registered stocks. Nestlé was rapidly followed by most large companies such as Sandoz, Ciba, etc. See Loderer and Jacobs (1995) for an analysis of the effect of the Nestlé equity restructuring.

[5] The Swiss Market Index is a continuously computed stock market index which is market value weighted and contains the 23 stocks having the largest market value and trading activity.

3. Data, Model Specification and Estimation Procedure

This chapter presents the data used in this study, discusses the choice of the instrumental variables and develops the stock index excess returns' model specification as well as the estimation procedure conducted to estimate its parameters.

3.1. DATA

We use the Swiss Banking Corporation General Index (hereafter SBC index) as the market portfolio proxy to test the risk premium effects induced by the three subsequent derivatives' listings. It is a broadly based stock market weighted index which comprises all traded bearer, registered and non-voting stocks traded on the Swiss stock market (approximately 400 stocks representing 99% of the Swiss market capitalization). The study relies on the SBC index values including the reinvested dividend stream as reported from Datastream. The subsequent tests all rely on weekly data reported each Wednesday over the period January 1988 until November 1992, that is based on a total of 253 weekly returns observations. The starting point for the estimations has been chosen as a compromise between the statistical need for a sufficiently long period of observations and the fact that we wanted to abstract from potential effects of the October 1987 crash.[6] Thus, the impact of the first stock option listing on May 19th, 1988 will have to be interpreted cautiously given the small number of prior observations (19) in the GARCH (1,1) model estimation. The relevance of the sample size and its anticipated impact on the results are further discussed in Section 3.3.

Also, it is worthwhile to emphasize that the choice of a weekly time interval has been motivated by the fact that Switzerland is a thinly traded stock market. Therefore, the estimation of a GARCH model based on daily returns would have been subject to estimation errors due to the lack of liquidity (leading to stale prices for some stocks or bid-ask quotes induced negative serial correlations in a large number of SBC index constituent stocks).

In Table I, the SBC index' returns and squared returns autocorrelations and partial correlations up to lag 12 suggest that, a priori, the choice of a GARCH (1,1) specification for the conditional variance of SBC index' weekly returns is indeed supported by the data. A couple of studies[7] have shown that it is indeed a valid representation for the Swiss stock market conditional variance dynamics.

Before turning to the description and estimation of the conditional asset pricing model used in this study, we should mention that the latter is based on the SBC index weekly excess returns computed with respect to the one month Euro Swiss francs interest rate (Euro CHF rate) adjusted for a weekly interval. The Financial Times Euro CHF rate series initially collected from Datastream over the reference

[6] Indeed, a GARCH specification cannot validly describe events such as the one associated with the October 1987 crash.

[7] See in particular Sabbatini (1996).

period has been double-checked with an independent data source provider (Datastream British Bankers Association series). Since the former presented unusual spreads (up to 8% of relative quote differences between the two Euro rate series) and was deemed by Datastream itself to be less reliable, we chose to rely on the second database, which only provides monthly Euro CHF rates series.

Following several studies in the United States (see in particular, Ferson and Harvey (1993)), Dumas and Solnik (1995) and De Santis and Gerard (1997, 1998)), we chose in the next section to represent the time-varying risk premia as an exponential function of a choice of instrumental variables. Five candidate variables were initially selected, that is the weekly change in the one month Euro CHF rate (Delta Euro), the slope of the yield curve as represented by the difference between the average yield-to-maturity of Swiss Government bonds maturing in more than 10 years and the one month Euro CHF rate (EYLT), the default spread, computed as the difference between the average yield-to-maturity on all Swiss Corporate bonds and the Long Term Government Bonds Datastream yield Index (DRYS), the lagged domestic SBC index excess weekly return (LDXR), and finally the lagged Financial Times and Standard & Poor's World Index[8] excess weekly return (LWXR) to account for the fact that Switzerland is an open economy. The summary statistics on the SBC index weekly excess returns and the candidate instrumental variables are all displayed in Table II. We finally describe the correlations between the set of instrumental variables in Table III. It is not surprising to observe that the most noticeable multicollinearity problems in the specification of the premium occur between both lagged indexes' returns and between each of the lagged index's return and the term spread series. Confirming recent empirical evidence from the term structure literature (see in particular Duffee (1998)), we also observe a negative correlation of -0.40 between the term spread and the default yield spread. Since the corporate yield spread series which should proxy for the time variation in the default risk in the economy is computed without any refinements with respect to the component bonds and their ratings, we decided to exclude the latter from the characterization of the time-varying market premium. The dividend yield or excess dividend yield over the one month Euro rate is often chosen as an instrumental variable since several studies (see in particular Kothari and Shanken (1996)) have shown its ability to track the time variation in expected real stock returns in the United States. Unfortunately, there is no reliable dividend yield index to be retained for the Swiss market which, coupled with the fact that Swiss companies tend to smooth their dividend policy, led us to refrain from using the dividend yield as an instrumental variable.

[8] Currency risk is neglected since this index is expressed in a synthetic currency representing the thirty included markets.

3.2. SPECIFICATIONS OF THE GARCH (1,1)-M SBC STOCK INDEX EXCESS RETURNS GENERATING MODELS

This section of the paper introduces the class of models chosen to describe the first two moments of the SBC index returns distribution. The most efficient model will then be used in Section 4 to assess the effects of derivatives listings on the unitary risk premium.

The general class of excess returns generating models retained in this study can be expressed as

$$\begin{aligned} r_{M,t} &= \lambda_0 + \lambda_t.\sigma_{M,t}^2 + \varepsilon_{M,t}, \\ \varepsilon_{M,t} &= \Im_{t-1} \sim i.d.\left(0; \sigma_{M,t}^2\right), \end{aligned} \tag{1}$$

where $r_{M,t}$ is the excess rate of return of the SBC index at date t, λ_t is the unitary risk premium and $\sigma_{M,t}^2$ is the variance of the SBC index. $\Im_{t-1}$ is the set of information available at date $t-1$. The justification of the intercept λ_0 is discussed later in this section.

Following Merton (1973), λ_t can also be viewed as the coefficient of relative risk aversion of the representative agent derived from his indirect utility function.[9] This coefficient is bound to be positive if the agent is risk averse. If no additional assumptions are made at this point, as mentioned in Section 1, nothing prevents λ_t from varying intertemporally. In Section 3.4, we tackle the relevance of constraining this coefficient to be constant over time. Following Harvey (1991) and Solnik (1993) for example, we will express λ_t as the exponential of a linear combination $k'.Z_t$, $k' = \left[k_{1,\ldots}, k_i\right]$, of a set of instrumental variables Z_t that also includes a constant term

$$\lambda_t = \exp\left(k'.Z_t\right). \tag{2}$$

Hence, in Equation (2), the exponential function precludes any negativity of the coefficient of relative risk aversion and the linear combination nests the constant case if the coefficients k_i are null for each instrument except for the coefficient of the constant term k_1. The set of instrumental variables contained initially a constant term and the five instrumental variables described in Section 3.1. Comparisons of likelihood performances[10] conducted on various subsets of Z_t led us to choose two instrumental variables, which jointly account for nearly all the variations in the premium. Thus, the constituents of the information set used in the empirical estimations conducted hereafter are a constant term, the weekly change in the one month Euro CHF rate (Delta Euro) and the lagged Financial Times/Standard & Poor's World Index excess weekly returns (LWXR).

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[9] Strictly speaking, this assertion holds when the difference in the variances of the wealth and consumption relative changes is negligible.

[10] See the discussion on the relevance of likelihood ratio test statistics in a quasi-maximum likelihood context in the next section.

Some authors (for instance, De Santis and Gerard (1997) in an international conditional CAPM setting) also specify a linear function in Equation (2). For the sake of comparison, we have estimated such a version of Equation (2). Since this specification yields similar results in term of likelihood and does not guarantee the positiveness of the coefficient of relative risk aversion, a fact that is difficult to sustain from a theoretical point of view, this specification was not deepened further.

Another issue to be considered is the presence of an intercept λ_0 in Equation (1). Market imperfections such as taxes or transaction costs[11] could lead to an intercept in Equation (1). It follows that the CAPM defined in a perfect market setting should theoretically imply an intercept λ_0 equal to zero since $r_{M,t}$ is defined as an excess return. However, for the sake of the econometric estimation, we will keep this intercept in the equation without altering the interpretation of the coefficient λ_t defined as the unitary risk premium and of the total market premium defined as the product $\lambda_t.\sigma^2_{M,t}$. In light of the stable transaction costs structure and fiscal policies prevailing in Switzerland during the sample period, we are assuming that λ_0 is constant over the sample.

If one removes the non-redundancy assumption, derivatives listings' effects, if any, could a priori alter either the conditional variance process of the market index $\sigma^2_{M,t}$ or the unitary risk premium λ_t or both parameters constituents of the total market premium. Section 3.3 first describes the process of the conditional variance of market excess returns, it then specifies the unrestricted version of the SBC Index excess returns model which allows derivatives listings to modify both the conditional variance process and the unitary risk premium at each derivatives' introduction and finally, the econometric procedure used to estimate the parameters.

3.3. METHODOLOGY AND ESTIMATION PROCEDURE

This section further describes the conditional variance of the market excess returns model used in this study and the econometric methodology applied to estimate its parameters and to test the different parameterizations of the index excess returns generating model.

The conditional second moments are modelled according to a GARCH (1,1) specification which allows the past conditional market variance and the past market innovation to appear in the current conditional variance equation. Bollerslev (1986) introduced GARCH processes, a generalization of the seminal paper of Engle (1982) on autoregressive conditionally heteroskedastic (ARCH) processes. The GARCH (1,1) variance process is defined as follows

$$\sigma^2_{M,t} = C^2 + a^2.\varepsilon^2_{M,t-1} + b^2.\sigma^2_{M,t-1}. \tag{3}$$

Looking at Equation (3), it is clear that $\sigma^2_{M,t}$ will be non-negative under very weak conditions. Moreover, if the GARCH parameters a and b verify the condition

[11] See Scruggs (1998) for a further discussion of this issue.

$a^2 + b^2 < 1$, then the process $\{\varepsilon_{M,t}\}$ is covariance stationary and, following Ding and Engle (1995), equation 3 can be restated as

$$\sigma^2_{M,t} = \sigma^2_0 \left(1 - a^2 - b^2\right) + a^2.\varepsilon^2_{M,t-1} + b^2.\sigma^2_{M,t-1} \tag{4}$$

where σ^2_0 is the unconditional variance of the residuals.[12]

GARCH-in-Mean models developed by Engle, Lilien and Robins (1987) combine a set of conditional mean equations with a process for conditional second moments. In this study, we combine Equations (1), (2) and (4) and introduce a set of dummy variables in the mean and the second moment equations that will take derivatives listing effects into account. Let $D_{i,t}$ be a dummy variable equal to 1 if $d_i \leqslant t < d_{i+1}$ where d_i is the date of the ith, derivative contract introduction and 0 elsewhere.[13] For the last introduction, the dummy variable is equal to one from 9th November 1990 (date d_3) to the end of the sample. Let k be a 3×1 vector of coefficients of the instrumental variables; l, n and p are 3×1 vectors of coefficients characterizing the unitary risk premium during the subperiods defined by the three dummy variables and the coefficients ζ_i capture level changes in the conditional variance process due to the ith derivative introduction.[14] These features yield the following general system

$$r_{M,t} = \hat{\lambda}_0 + \exp\left(\left(k' + D_{1,t}.l' + D_{2,t}.n' + D_{3,t}.p'\right).Z_t\right).\hat{\sigma}^2_{M,t} + \hat{\varepsilon}_{M,t}, \tag{5}$$

where

$$\begin{aligned} \hat{\sigma}^2_{M,t} &= \hat{\sigma}^2_0 \left(1 - a^2 - b^2\right) + a^2.\hat{\varepsilon}^2_{M,t-1} + b^2.\hat{\sigma}^2_{M,t-1} + \zeta_1.D_{1,t} \\ &\quad + \zeta_2.D_{2,t} + \zeta_3.D_{3,t} \\ \hat{\varepsilon}_{M,t} &= \Im_{t-1} \sim i.d.\left(0; \hat{\sigma}^2_{M,t}\right). \end{aligned}$$

The specification of the dummy variables chosen does not capture transitory but only structural shocks on the unitary risk premium and on the conditional variance process. One should also notice that the second dummy variable will cumulate the effect of the first two derivatives listings and, in the same manner, the third dummy variable will capture the cumulative effects of the three derivatives' introductions. One may conjecture whether such a specification of the excess returns generating model can separate the pure listing effects from the time-induced variations in the unitary risk premium. However the structure of the dummy variables used in this

[12] An in-depth theoretical treatment of these conditions extended to multivariate processes can be found in Engle and Kroner (1995).

[13] The dummies respectively refer to:

d_1: May 19th, 1988 (individual stocks options listing date)

d_2: December 7th, 1988 (SMI options listing date)

d_3: November 9th, 1990 (SMI futures listing date)

[14] Following the empirical results in Bruand and Gibson (1998) on the significant effects of newly listed derivatives on the unconditional and conditional levels of the Index returns' variance, we only introduce dummy variables that capture level changes in the conditional variance process.

study prevents us from capturing 'pure' time variations in the instrumental variables' coefficients, at least as far as short cycle time variations of these coefficients are concerned. Indeed, the first dummy variable takes positive values over more than six consecutive months and the two subsequent dummies over approximately two consecutive years. Hence, any time variation of duration shorter than these respective lengths will not bias the estimated coefficients associated with the listing effects. We might however be concerned with time variations whose cycles are approximately twice as long as the length of the positive values of the dummies. For instance, the third dummy variable's coefficient may be sensitive to very long-term time variations of about four years. One might consider introducing a slowly mean-reverting instrumental variable that would impact the agent preferences to take these potential long-term time variations of the unitary risk premium into account. The exact specification of such an instrumental variable and its associated impact on the unitary risk premium, if any, are very interesting. They are however beyond the scope of the present study and are left for future research.

Before presenting the estimation results in Section 3.4, let us comment on two estimation problems, the first arising from the clustering of the listing dates and the second related to the small sample size. The latter is due to the fact that we only have 19 weekly observations before the individual stocks options listing, as explained in Section 3.1. The clustering effect is induced by the fact that there are only 28 additional observations between the individual stocks and SMI options listing dates. This clustering might affect the magnitude and the significance of the first two derivatives listings' effects on the conditional variance process and/or on the unitary risk premium. The meaningful test for individual and joint structural listings' effects would have necessitated a prior sample period estimation of the time-varying risk premium. We however felt that such an exercise was not very informative given that the prior period was characterized by other major economic events such as the stock market crash of October 87 which may have also influenced agents' preferences. Including observations from this unstable period would thus have added a lot of noise to the estimation.

The choice of the specific starting date leads to only nineteen observations prior to the first derivatives' introduction. This might seem insufficient to insure a proper econometric estimation of the coefficients. However, given that the dummy variables manage to fully seize the derivatives listings effects, these dummy variables will separate the effects of the three derivatives listings from the underlying market excess returns process over the entire sample. In other words, the vector k underpinning the model is estimated over the entire sample period, net of the structural shocks due to the three consecutive derivatives' listings.

The testing procedure for the derivatives listings' effects is based on the average difference between the net time-varying unitary risk premium $\left(\exp\left(k'.Z_t\right)\right)$ and the actual unitary risk premium including the derivatives listings $\left(\exp\left(\left(k'+p'\right).Z_t\right)\right)$ over the period following the last introduction (observations 149 through 253). This testing procedure is thus fairly immune to the biases due to the small number of observations prior to the first listing and due to the clustering of the first two derivatives introductions.

Since the unconditional variance σ_0^2 is not directly observable, we use the iterative method proposed by De Santis and Gerard (1997b). A consistent estimate is obtained by setting σ_0^2 equal to the sample variance of the SBC Index excess returns in the first iteration and then updating it, at each iteration, using the variance of the estimated residuals.[15]

We use the quasi-maximum likelihood (QML) method developed by Bollerslev and Wooldridge (1992) to estimate the parameters of the system in Equation (5) and to conduct hypothesis testing. The QML estimator (QMLE) is obtained by maximizing a normal log-likelihood even if the normality assumption is violated.[16] Under regularity conditions,[17] QMLE of the parameter is still consistent and has a limiting normal distribution with a covariance matrix equal to $H_T^{-1}.G_T.H_T^{-1}$ where H_T is minus the expected Hessian and G_T is the outer product of the gradient. This matrix is a consistent estimator of the White (1982) robust covariance matrix. We see that, under normality, the covariance matrix of QMLE reduces to either G_T^{-1} or H_T^{-1}. One should mention that this method is reliable provided that the first two moments are correctly specified (see Bollerslev and Wooldridge (1992) for the theoretical treatment).

QMLE is the vector $\hat{\theta}_T$ that maximizes the following conditional quasi log-likelihood

$$L_T(\theta) = -\frac{T}{2}\ln 2\pi - \frac{1}{2}\sum_{t=1}^{T}\ln\sigma_{M,t}^2(\theta) - \frac{1}{2}\sum_{t=1}^{T}\frac{\varepsilon_{M,t}^2(\theta)}{\sigma_{M,t}^2(\theta)}. \tag{6}$$

We used the BHHH (Berndt, Hall, Hall and Hausmann (1974)) optimization algorithm to maximize Equation (6) with the initial condition $\hat{\varepsilon}_{M,0} = 0$.[18]

The likelihood ratio test is not asymptotically chi-square distributed when the probability distribution of errors is not normal, which might be the case in the present quasi-maximum likelihood context. Thus, the relative difference between two likelihoods remains informative but it is difficult to test its statistical significance with the likelihood ratio test. Wald and Lagrange Multiplier tests are robust in the QML context. In the following analysis, we base our inferences on robust Wald tests but we nevertheless provide some likelihood ratio test statistics based on the chi-square distribution for informative purposes only. We emphasize that

15 We do not estimate σ_0^2 as a free parameter in the log-likelihood for computational reasons.

16 We thus assume that the residuals are conditionally normally distributed in the conditional mean equation.

17 Interested readers should look at Conditions A.1, p. 167 in Bollerslev and Wooldridge (1995).

18 The BHHH algorithm is an iterative maximization method defined as follows: $\theta^{i+1} = \theta^i + \tau_i \cdot (S'S)^{-1} S'\iota$, where i is the iteration number, $[S]_{tk} = \partial L_t/\partial\theta_k$ is the score, k is the number of coefficients, τ_i is the step length found by a linear search and ι is a vector of ones. The increment is thus found by regressing a vector of $1's$ on the scores.

these likelihood ratio test statistics might be inappropriate if the true distribution of errors departs from normality.

We compute robust t-statistics and tests on the differences of means to measure the significance of the derivatives listings' effects on the unitary market price of risk.[19]

3.4. CHOICE OF THE SBC INDEX EXCESS RETURNS MODEL AND TIME-VARYING RISK PREMIUM EFFECTS

This section describes the estimation results that allow us to estimate the significance and magnitude of derivatives' listings effects on the SBC Index excess returns. We report in Table IV the parameterization of model I which corresponds to Equation (5) as well as the different restrictions which define the six other models (from II to VII). In the absence of any priors, in model I, the three listings can affect the level and the slope coefficients of the risk premium. However, for the sake of parsimony, we rely on former empirical results in Gibson and Bruand (1998), in only allowing listings' effects to modify the constant of the conditional variance in the model I. The number of parameters and the respective log-likelihoods for the seven models are also provided in this table. The number of parameters implied by their parameterizations can be quite large, as illustrated by the specification of model I. Thus, parsimony is of great concern.

We first test the usefulness of including a constant term λ_0 in the mean equation. In Table 4, we see that model I (no restriction) has a conditional log-likelihood equal to -505.85. Model II (intercept constrained to zero, $\lambda_0 = 0$) leads to a log-likelihood equal to -507.34. While λ_0 is not statistically significant (robust t-stat of -0.99) in Model I, the difference of likelihoods between Model I and II suggests that the constant has some role.[20] Given that this coefficient might capture market imperfections (such as taxes, for example), we prefer not to take the risk of biasing the remaining estimated coefficients and thus the restriction $\lambda_0 = 0$ will not be imposed in the following estimation procedures.

We are not aware of any empirical study, which allows the unitary risk premium to be affected by derivatives listings. Hence, in a first stage, we wish to determine whether derivatives listings affect the conditional variance process only or the unitary risk premium only, or both. We therefore estimate models III and IV (see specifications in Table IV). Model III allows only the constant of the GARCH process to be modified at introduction dates and imposes a constant unitary risk premium. Model IV is featured by a constant GARCH process over the estimation period and allows each coefficient k_i to be affected by each introduction. Model

[19] In the present setting, asymptotic least-squares can also be used for the testing procedure. This method is robust if the robust covariance matrix of the estimators is used.

[20] As discussed in the previous section, the likelihood ratio test statistics are not asymptotically chi-square distributed if the true distribution of errors is not normal. Hence, we cannot rely on this test to make a decision.

I nests these two models since it allows both the constant of the GARCH process and all the coefficients k_i appearing in the specification of the risk premium to be modified independently of the other at any of the three introductions dates. The Wald statistic of the restrictions imposed in model IV (shown in Table IV) is equal to 2,275. Thus, it seems that the derivatives listings affect the unitary market risk premium only but model V will refine this conclusion as shown subsequently.

Testing for derivatives listings' effects on the unitary premium is a rather delicate task. Wald tests on model I provided in Table V assert that we cannot reject that there are listing effects on the unitary risk premium (p-value = 0.0667) and, more generally, that the unitary risk premium is time-varying due to listing effects and to the changes in the instrumental variables (p-value = 0.0095). Without anticipating again on results given in the following paragraph, Wald tests imply a mixed effect of derivatives listings' on both the GARCH process and on the specification of the unitary risk premium. This conclusion is furthermore supported by the Wald statistic reported in Table V which tests the hypothesis of null listing effect over the sample period (p-value = 0.0657).

Finally, there might be some redundancy among all dummies retained in model I: we allow each coefficient k_i to vary at each of the three derivatives introductions, hence yielding nine dummies to describe changes in the process of λ_t through the sample period and three additional dummies for changes in the level of the conditional variance process. Some of these dummies might only add degrees of freedom in test statistics but no additional information in the log-likelihoods themselves. Thus, we performed wide range of Wald tests[21] on model I to see if we can constrain some of its coefficients to zero without loosing significant explanatory power. For example, in one of these Wald tests, we restrict the coefficients of the dummies on the two instrumental variables to zero to see if the derivatives listings only impact the unitary risk premium through the level coefficients. These tests helped us to identify six redundant parameters in the specification of model I that lead us to empirically select model V (see Table VI for the specification and the parameters estimates of model V). These six parameters are redundant since constraining them to zero does not lead to a significant loss of explanatory power. In other words, the constraint is not statistically binding. Exclusion of these six irrelevant parameters increases the precision of the final estimates. In Table V, the Wald test statistic for the restrictions imposed in model V (12 parameters) is equal to 1.9273 with a p-value of 0.9263. Hence, it confirms the irrelevance of the six excluded parameters.

The choice of model V also confirms that derivatives listings imply a mixed effect since two of the three dummies associated with the conditional variance process remain and since seven of the twelve parameters used in the specification of the unitary risk premium in model I are also present in this restricted model. Examining the restrictions imposed in model V, we see that most of the effect on the unitary risk premium is captured through modifications of its sensitivity to the

[21] These tests are available upon request.

second factor, the lagged world index excess return (LWXR), since all three dummies on this instrument are significant (in Equation (5), l_3, n_3 and p_3 are therefore significantly different from zero).

Looking at the parameter estimates of model V in Table VI, we can see that the condition on GARCH parameters for covariance stationarity of the residuals is satisfied since $a^2 + b^2 < 1$. We report both the value of a and b and their squared value a^2 and b^2 respectively for ease of comparison. T-stats for the squared values are computed with the delta method. The statistical significance of both a^2 and b^2 is well established with one-sided p-values of 0.0365 and 0.000 respectively. The persistence is equal to 0.8336 giving a half-life of the return shocks of approximately one month (3.81 weeks).

Considering the dummies related to the three derivatives launching effects in model V, the mean of the unitary risk premium (λ_t) over the sample period given in Table VII is about 0.1447 implying a risk aversion coefficient on the Swiss market of 14.47.[22] This is a fairly high level of the coefficient of risk aversion but looking at the graph displaying the evolution of the unitary risk premium λ_t given in Figure 1, one can note that the unitary risk premium is only very high between the first and the second derivatives listings. The mean of λ_t excluding this period is about 0.0987 which seems more plausible. The explanation for the high degree of the unitary risk premium between May and December 1988 is still not clear.[23]

4. The Effects of Derivatives Listings on the Unitary Market Risk Premium

As already mentioned in the introduction, previous studies on the role of financial innovation ignored the joint impact of derivatives listings on the market portfolio variance and risk premium specifications. As a result, and as shown in Section 3, they are not powerful enough with respect to their main conclusion which asserts that only the stock and/or stock index variance are indeed significantly affected by successive listings. However, once the joint time-varying unitary risk premium and GARCH (1,1) specifications of the conditional market excess returns model is considered and tested for alternative nested models with respect to the potential

[22] We have to multiply the mean lambda by 100 since estimations were conducted with relative figures. As already mentioned, this definition of the relative risk aversion coefficient only applies when the wedge between the variances of the aggregate consumption and wealth relative changes is small.

[23] One could conjecture that the modification of the ownership structure, discussed in Section 2, that occured at the same time in Switzerland might explain this increase in the unitary risk premium. Indeed, restrictions on registered shares ownership transfer were reduced by major firms in November 1988. Bearer shares used to traded at a premium of 20% to 40% over restricted registered ones before November 1988. Gardiol, Gibson and Tuchschmid (1997) study this "bearer" shares' premium in Switzerland over the period 1980–1992 and show that, while persistent and time-varying, this premium cannot be arbitraged. This suggests thus that both bearer and registered stocks were imperfect substitutes. However, this reduction in ownership restrictions boosted the liquidity of the concerned shares and thus should have led, if anything, to a lower unitary risk premium.

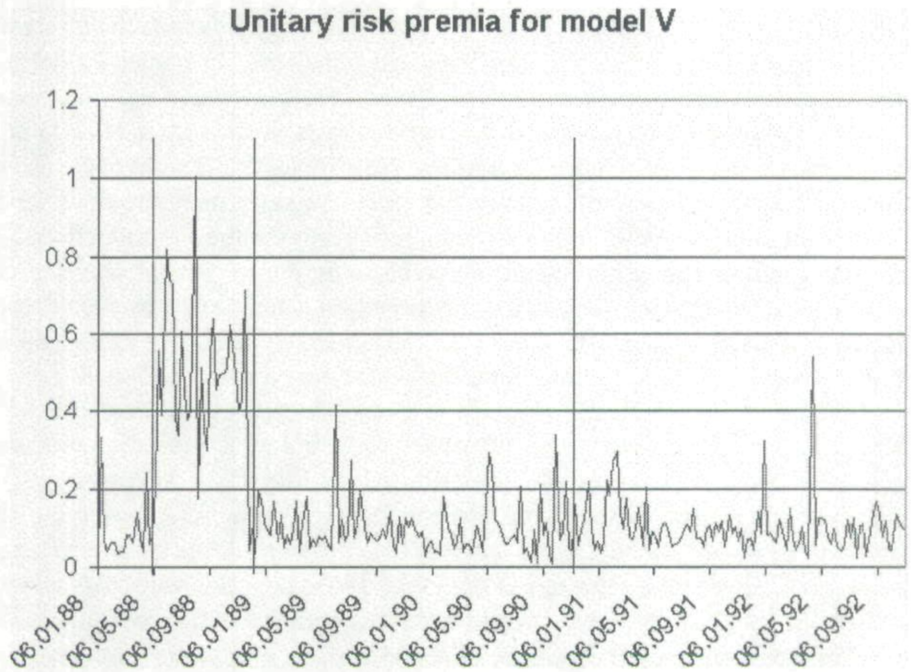

Figure 1. Time-varying unitary market risk premium on the Swiss market. This graph shows the evolution of the time-varying unitary risk premium λ_t computed with model V over the sample period (6th January 1988 to 11th November 1992). Parameterization and QML estimators of model V are given in Table VI. The time-varying unitary risk premium plotted here includes the effects of the derivatives listings. Vertical lines identify the three derivatives listings which take place on the 19th May 1988, the 7th December 1988 and the 9th November 1990 respectively.

impacts of the three listings, these conclusions can be refined. In Table V, the most parsimonious nested model V retained on the basis of its likelihood and several Wald tests with respect to the model comparison discussed in Section 3.2, displays two dummies (for the first and the third listings) on the conditional variance specification, while it displays different channelling mechanisms for the structural changes in the unitary risk premium subsequent to the three listings. More specifically, we see that the model allows for a significant change in the constant term and in the second instrumental variable (LWXR) of the risk premium specification when individual stock options were first introduced. It captures an effect with respect to both instrumental variables (Delta Euro and LWXR) at the subsequent introduction of the SMI index options and, allows the second instrumental variable (LWXR) to capture any significant effects related to the final introduction of the SMI futures contracts.

Notice that the observed variations of the total risk premium in model V after each successive listing are consistent with the gradual market completion hypothesis invoked by Detemple and Jorion (1990). These authors assert that, as financial innovation extends, the market reaches a stage in which the effect of any additional derivative's introduction is indeed redundant. In model V, the gradual market completion hypothesis holds with respect to the level of the stock market's conditional variance effects since the two dummies of the GARCH (1,1) process are negative but gradually less significant (see Table VI). The hypothesis is also sustained with respect to the unitary risk premium since there are fewer and less significant dummies when we move from the first to the third introduction.

The gradual market completion hypothesis may also explain why the coefficients of the constant dummy in the risk premium is unaffected by the second and the third listings although the latter result may also be due to the clustering of the two first listings which occurred within less than seven months.[24]

Finally, another interesting feature is that while the second instrumental variable, i.e. the lagged excess return of the World Index, is indeed significant and captures effects at the three derivatives listings, the Euro CHF rate sensitivity of the risk premium is not affected by the first and the third listings. Moreover, the second options' listing effect on the risk premium sensitivity with respect to changes in the Euro CHF interest rate (n_2) is not clear. Indeed, while Wald tests conducted in Section 3.4 do not allow us to constrain this parameter to zero in model V, the t-stat of its estimate is not significant. We checked if the latter phenomenon was not artificially induced by two key modifications in the Swiss National Bank monetary policy, that is, the introduction of the SIC (Swiss Interbank Clearing) system and the necessary changes in the Banks monthly liquidity prescriptions (corresponding to a reduction from 6 to 2.5% of their short term liabilities and 20% of their savings and deposits holdings) but these modifications were both enforced on the 1st of January 1988. There was no other major structural change in the Swiss National Bank's monetary policy during the period examined.[25]

We now turn to the main empirical issue, namely have the three derivatives listings increased or decreased the unitary market risk premium? As stated in the introduction, one would be inclined to think that with a broadening of the market's investment opportunity set, gradual market completion should lower the unitary risk premium in order to reflect the fact that risk-averse investors are now, all other things being equal, better off in terms of satisfying their risk management demands and face implicitly lower liquidity constraints (see Detemple and Serrat (1998)). They gain more flexibility and, in addition, they can benefit from the two other economic functions, that is asset allocation and price discovery provided by derivatives to market participants who could only trade in the cash market. As already said, this is mainly a conjecture since there is no general equilibrium model cast in an intertemporal setting which looks at the completion effects of derivatives launchings and at their impact on a time-varying market risk premium. In addition,

24 As mentioned in Section 3, any analysis of the separate effects of those two events needs to be interpreted cautiously due to their clustering.

25 See Nilles (1992) for a complete treatment of the Swiss National Bank's monetary policy over the sample period.

we cannot exclude that the conclusions stemming from a one-period asset pricing model cast in an incomplete market (such as Detemple et al. (1990, 1991)) may not necessarily carry through in an intertemporal setting.

We find that overall, the three listings jointly reduced the unitary time-varying risk premium by 5.53%. The reduction in the unitary risk premium is analyzed in Table VII in which we display the average level of the premium with and without effects induced by the three derivatives listings. The mean of the "base" unitary risk premium[26] which does not incorporate the listings dummies ($= \exp\left(k'.Z_t\right)$) amounts to 0.1066 over the period following the third introduction while the mean of the unitary risk premium including the effects of the three listings is equal to 0.1007 over the same period. This reduction is significant since the robust t-stat of the dummy after the third introduction (p_3) is equal to 2.10. However, one should note that, based on a simple test of the difference of means of the 105 observations in the two cases, this significance disappears. Figure 1 displays the evolution of the unitary risk premium given by model V.

As shown in Table VII, the first introduction increased the risk premium since the average of the linear combination $D_{1,t}.l'.Z_{t-1}$ is equal to 1.8244 (see Table VII), hence giving a multiplicative coefficient of the "base" risk premium of $\exp(1.8244)$ or 6.199. This explanation for the high level of the risk premium, which appears in Figure 1 during the period, May–December 1988 must however be treated with skepticism.[27] The cumulative effect of the first two introductions ($= D_{2,t}.n'.Z_{t-1}$) seems to decrease the "base" risk premium by a multiplicative coefficient of 0.857 ($= \exp(-0.1545)$). The three introduction cumulatively yielded an average multiplicative coefficient of 0.975 ($= \exp(-0.025)$). Jensen's inequality explains the difference between this latter multiplicative coefficient and the overall effect fostered in the previous paragraph.

4.1. RELEVANCE OF THE UNITARY RISK PREMIUM SPECIFICATION AND THE SMOOTHING EFFECT

The sensitivity of the risk premium obtained in this and previous studies on time-varying asset pricing models to the specification and the choice of the instrumental variables is rarely tackled. Indeed, the number and the choice of the instrumental variables used to capture the temporal variations in the unitary risk premium are most of the time selected on an ad hoc basis. Adding an instrumental variable clearly increases the information set but it can also have a smoothening effect by reducing the impact of the specific shocks associated with other series included in the set. In order to illustrate our point, we repeat the derivatives listings' effect test-

[26] I.e. the vector of coefficient k that prevails over the entire sample.

[27] We checked for any significant economic, politic or institutional events that occured during that period. None of them seems to justify such an important increase in the level of the unitary risk premium. On the other hand, the launching of individual stock options does not per se suffice to explain the magnitude of the observed increase of the unitary risk premium.

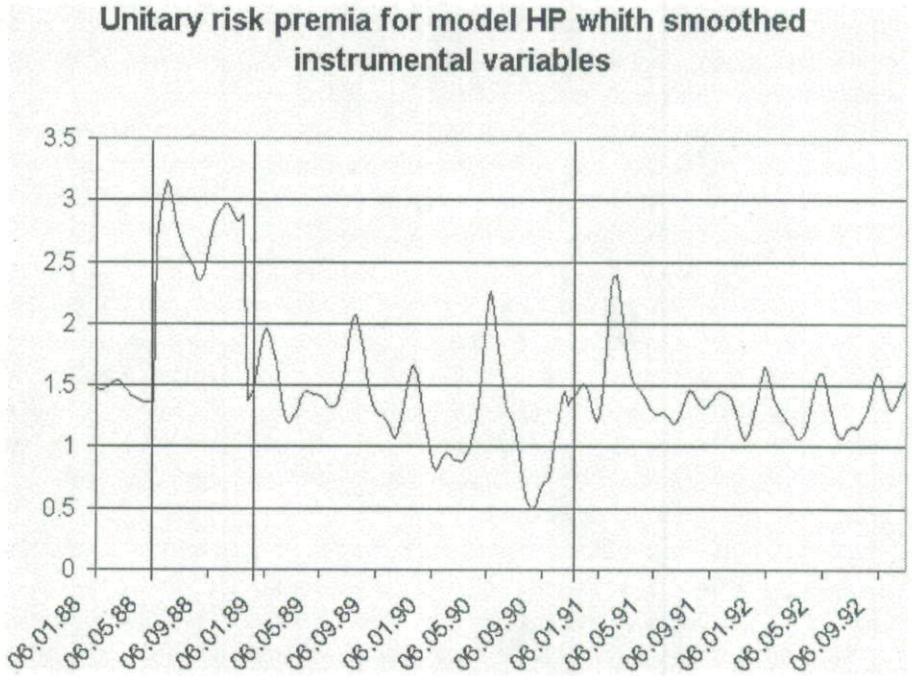

Figure 2. Time-varying unitary market risk premium on the Swiss market with smoothed instrumental variables. This graph shows the evolution of the time-varying unitary risk premium λ_t computed with model HP over the sample period (6th January 1988 to 11th November 1992). Instrumental variables are Hodrick-Prescott filtered in model HP before the estimation procedure. Parameterization and QML estimators of model HP are given in Table VI. The time-varying unitary risk premium plotted here includes the effects of the derivatives listings. Vertical lines identify the three derivatives listings which take place on the 19th May 1988, the 7th December 1988 and the 9th November 1990 respectively.

ing procedure with smoothened instrumental variables. Since this is not a standard procedure in the conditional CAPM literature, we choose a Hodrick-Prescott filter of 10 to smooth the two series of instrumental variables and conduct the estimation procedure with the same specification as in model V. The estimated coefficients of this "smoothed" model, named model HP are provided in Table VI. The main results are that (i) the log-likelihood is higher by approximately 20 points, (ii) the unitary risk premium is multiplied by a factor of 10 on average[28] and (iii) the conditional variance of the SBC index excess returns is less persistent and lower on average by 17 percent. The representation of the smoothed unitary risk premium (according to model HP) is provided in Figure 2.

These results may partially explain why, when we look at the literature on time-varying risk premia, there is no consensus on the optimal number of instrumental variables nor clear evidence on which ones should constitute the relevant inform-

[28] The average unitary risk premium over the entire sample for the model with smoothed instrumental variables is equal to 1.5175 against 0.1447 for model V.

ation set. Indeed, adding an instrumental variable in the information set (Z_t) may not add any additional information but help to smooth the functional form of the equation describing the unitary risk premium (here Equation (2)). This smoothing effect will of course depend on the volatilities and covariances among the set of instruments; the less correlated the instruments, the better the smoothing. Hence, some authors (see e.g., Jagannathan and Wang (1996)) use only one instrumental variable (the default spread) while some incorporate more than five series in the information set.[29] The derivatives listings effects captured in the smoothed setting are consistent with those found in the previous Section as far as their sign is concerned. However, their magnitude and significance differ in the HP model setting. It turns out that the determination of the appropriate smoothing method and its impact on the choice of the instrumental variables set play an important role not only for the measurement of derivatives listing effects but more generally for the assessment of the process driving the time-varying risk premium. This topic should hopefully stimulate further interest and research to enhance our understanding of time-varying preferences within intertemporal asset pricing models.

5. Conclusion

The empirical results highlighted in this study are certainly preliminary but nevertheless emphasize an important point, namely that the launching of derivatives significantly reduces the unitary risk premium, thus contributing to the literature on the time-varying property of the latter. This observation poses a challenge if one wishes to reconcile it with most asset pricing models which – even if cast in an intertemporal framework – still treat derivatives as purely redundant securities. According to the preliminary results obtained in this study, the stock market conditional variance as well as the components of the unitary time-varying risk premium and thus, equilibrium cash prices, are indeed affected by the gradual market completion process, especially at its early stages. Our findings are consistent with the theoretical lines of Detemple and Serrat (1998) if one is willing to consider the listing of derivatives as a market completion mechanism which lowers agents' liquidity constraints and thus in turn lowers the market Sharpe ratio. The significance of the overall reduction by 5.53% of the unitary risk premium observed on the Swiss stock market subsequent to the three listings must however be interpreted with caution. Indeed, this result rests on the validity of the joint hypothesis of a time-varying unitary risk premium GARCH (1,1)-M specification of the index excess returns coupled with the derivatives launching redundancy used in the empirical tests. Furthermore, the time period over which the study was conducted is fairly short. In light of these restrictions, the evidence suggests that for the Swiss stock market, the unitary and total time-varying risk premia declined

[29] Bossaerts and Hillion (1999) compare several selection criteria in a linear model setting with varying numbers of predictors. The 'best' set of predictors seems to be country as well as criteria-dependent.

significantly after the three derivatives were introduced. It would be interesting to conduct similar tests in other European, Asian and North American stock markets in order to see whether and to what extend the effect on the unitary and total market risk premia are related to institutional, domestic or international stock markets' characteristics. In the same vein, this study could be extended in order to assess the regulatory, welfare and economic policy implications of the launching of derivatives in LDC countries that are still at an early stage of their market completion process.

From a methodological perspective, two areas of research seem to be of particular interest. First, the joint specification of the time-varying unitary risk premium and conditional variance processes in most intertemporal asset pricing models deserves more consistency across theoretical and econometric modelling. This observation certainly applies when one studies the effects of derivatives launchings but it also extends to event studies focusing on the role of political, macroeconomic or regulatory changes – such as monetary policy changes, the introduction of the Euro, etc. – on agents' preferences. Secondly, we emphasized in this study how important the specification of the functional form of the unitary risk premium appears to be when testing for derivatives' listings effects. Surprisingly, most of the empirical literature has so far provided little consideration to the rather delicate questions of how one should choose the number and the characteristics of the instrumental variables in intertemporal asset pricing models and further specify the functional form of the time-varying risk premium. These issues clearly deserve further empirical investigations especially when one implements or compares the validity of alternative intertemporal asset pricing models.

Appendix

Table I. Correlogram of stock market excess returns and squared stock market excess returns

	Market excess returns				Squared market excess returns			
Lags	AC	PAC	Q-stat	Prob	AC	PAC	Q-stat	Prob
1	0.092	0.092	2.17	0.141	−0.007	−0.007	0.01	0.908
2	0.152	0.145	8.12	0.017	0.304	0.304	23.76	0.000
3	0.018	−0.007	8.21	0.042	−0.029	−0.028	23.97	0.000
4	0.022	−0.001	8.33	0.080	0.025	−0.075	24.13	0.000
5	0.005	0.001	8.34	0.138	0.202	0.243	34.73	0.000
6	−0.018	−0.022	8.42	0.209	0.078	0.103	36.33	0.000
7	0.074	0.079	9.86	0.197	0.111	−0.036	39.57	0.000
8	−0.065	−0.074	10.97	0.203	−0.012	−0.049	39.61	0.000
9	0.042	0.033	11.44	0.247	−0.021	−0.026	39.73	0.000
10	−0.017	−0.003	11.52	0.319	0.036	0.024	40.08	0.000
11	−0.115	−0.131	15.06	0.180	0.033	0.019	40.37	0.000
12	−0.114	−0.093	18.53	0.101	0.059	0.007	41.30	0.000

Statistics refer to weekly excess returns of the Swiss Bank Corporation General Index which covers all traded shares on the Swiss market. The weekly excess returns are obtained by subtracting the one month Euro Swiss francs interest rate expressed on a weekly basis. The market index includes reinvested dividends as computed by Datastream. The sample period of 253 observations is from January 6th, 1988 to November 11th, 1992. AC is the autocorrelation of the series at each lag k. PAC is the partial autocorrelation of the series at each lag k. Q-stat and Prob are the Ljung-Box Q-statistics and their p-values. The Q-statistic at lag k is a test statistic for the null hypothesis that there is no autocorrelation up to order k.

Table II. Descriptive statistics of the market index and the instrumental variables

	SBC Index	Delta Euro	DRYS	EYLT	LDXR	LWXR
Mean	0.021	0.014	0.010	−0.026	0.035	−0.004
Maximum	4.931	0.687	0.027	0.061	4.931	5.769
Minimum	−8.579	−1.187	0.001	−0.082	−8.579	−7.379
Std. Dev.	1.933	0.267	0.003	0.031	1.944	1.734
Skewness	−0.853	−0.600	0.621	1.155	−0.839	−0.602
Kurtosis	5.736	4.702	6.513	3.557	5.661	5.046
Jarque-Bera	109.56	45.73	146.38	59.51	104.35	59.46
p-value	0.00	0.00	0.00	0.00	0.00	0.00

Swiss Bank Corporation General Index (SBC index) is the market index proxy as described in Table 1. DELTA EURO is the first difference of the Euro CHF one month interest rate series from the British Bankers Association; it serves as risk free interest rate to compute all excess returns. DRYS is the default risk yield spread computed as the difference between average yield-to-maturities of all corporate bonds traded on the Swiss market and of long term (over 10 years) government bonds. EYLT is the excess yield of long term (over 10 years) Swiss government bonds over the Euro CHF one month rate. LDXR is the lagged (1 period) excess return of the domestic stock market index (SBC index) over the Euro CHF one month rate. LWXR is the lagged excess return of the Financial Times/Standard & Poor's World Stock Market Index over the Euro CHF one month rate. Both domestic and world stock market indexes are dividend reinvested. All series are provided by Datastream. The sample period of 253 observations is from January 6th , 1988 to November 11th, 1992. Kurtosis is equal to 3 for the normal distribution. Small p-values of the Jarque-Bera test lead to reject the null hypothesis of normal distribution.

Table III. Correlation matrix of the instruments

	DRYS	EYLT	LDXR	LWXR
Delta Euro	−0.066	−0.077	−0.092	0.035
DRYS		−0.404	−0.076	−0.041
EYLT			0.123	0.166
LDXR				0.705

This table gives the sample correlation matrix of potential instrumental variables. The two variables that were included in the information set for the estimations of models I, II, V and VI are : the first differenced Euro CHF one month interest rate series (DELTA EURO) and the lagged excess return of the Financial Times/Standard & Poor's World Stock Market Index over the Euro CHF one month rate (LWXR). DELTA EURO is the first differenced Euro CHF one month interest rate series from the British Bankers Association. The one month Euro rate also plays the role of the risk free interest rate to compute all excess returns. DRYS is the default risk yield spread computed as the difference between average yield-to-maturities of all corporate bonds traded on the Swiss market and of long term (over 10 years) government bonds. EYLT is the excess yield of long term (over 10 years) Swiss government bonds over the Euro CHF one month rate. LDXR is the lagged (1 period) excess return of the domestic stock market index (SBC index) over the Euro CHF one month rate. LWXR is the lagged excess return of the Financial Times/Standard & Poor's World Stock Market Index over the Euro CHF one month rate. Both domestic and world stock market indexes are dividend reinvested.

All series are provided by Datastream. The sample period of 253 weekly observations is from January 6th, 1988 to November 11th, 1992.

Table IV. Alternative model specification of the SBC index excess returns

Model	Restrictions	Param.	Log	Remarks
I	No restriction	18	−505.85	
II	$\lambda_0 = 0$	17	−507.35	No intercept in the mean equation
III	$k' = (k_1\ 0\ 0)$ $l = n = p = 0$	7	−512.53	Introductions are restricted to affect the variance process only. Constant unitary risk premium
IV	$\zeta_i = 0, \forall i$	15	−511.41	Introductions affect the unitary market risk premium only.
V	$\zeta_2 = 0$ $k' = (k_1\ 0\ k_3)$ $l' = (l_1\ 0\ l_3)$ $n' = (0\ n_2\ n_3)$ $p' = (0\ 0\ p_3)$	12	−507.74	Introductions affect both the variance process and some coefficients of the unitary risk premium.
VI	$k' = (k_1\ k_2\ k_3)$ $l = n = p = 0$ $\zeta_i = 0, \forall i$	6	−514.91	Time-varying risk premium and no introduction effect.
VII	$k' = (k_1\ 0\ 0)$ $l = n = p = 0$ $\zeta_i = 0, \forall i$	4	−516.78	Constant risk premium with no introduction effects.

The most general parameterization is provided by model I with, in the mean equation (A.1), $r_{M,t}$ the excess return of the SBC stock market index over the Euro CHF 1 month at time t, λ_t the unitary market premium function at time t (A.2) and, in the variance equation (A.3), $\sigma^2_{M,t}$ the conditional variance of the SBC index at time t. These three equations relate exactly to the system in Equation (5) in text. The set of instruments Z_t that explain the variations of the unitary market premium include a constant, the first difference of the Euro CHF 1 month rate (DELTA EURO) and the lagged excess return of the world index over the Euro CHF 1 month (LWXR). k, l, n and p are hence 3×1 vectors. The variance process is a GARCH (1,1) which enables each derivatives listing to modify the constant of the GARCH process through ζ_i. D_1 is a dummy which equals 1 between the first and the second derivative introductions (i.e. between the 19th May 1988 and the 7th December 1988) and 0 elsewhere. D_2 is a dummy which equals 1 between the second and the third derivative introductions (between the 7th December 1988 and the 9th November 1990). D_3 is a dummy which equals 1 after the third derivative introduction (9th November 1990).

Models II to VI are constrained versions of model I. Model II is the only model with no intercept in the mean Equation (A.1). Model III allows the parameters of the variance process to react at derivatives listings only while model IV has a constant GARCH process for the variance and a time-varying unitary risk premia which captures potential listing effects. Model V enables both types of effects. Models VI and VII do not incorporate any derivatives listings effects.

$$r_{M,t} = \hat{\lambda}_0 + \lambda_t.\hat{\sigma}^2_{M,t} + \hat{\varepsilon}_{M,t}, \tag{A.1}$$

$$\lambda_t = \exp\left[\left(k' + D_{1,t}.l' + D_{2,t}.n' + D_{3,t}.p'\right).Z_t\right], \tag{A.2}$$

$$\begin{aligned}\hat{\sigma}^2_{M,t} \;=\; & \hat{\sigma}^2_0\left(1 - a^2 - b^2\right) + a^2.\hat{\varepsilon}^2_{M,t-1} + b^2.\hat{\sigma}^2_{M,t-1} + \zeta_1.D_{1,t} \\ & + \zeta_2.D_{2,t} + \zeta_3.D_{3,t},\end{aligned} \tag{A.3}$$

$$\hat{\varepsilon}_{M,t} = \Im_{t-1} \sim i.d.\left(0; \hat{\sigma}^2_{M,t}\right).$$

N.B.: For checking purposes models VI and VII were estimated. They do not feature any derivatives listing.

Table V. Likelihood ratio and Wald tests for derivatives non-redundancy

Likelihood ratio test for the intercept[30]		
Is the intercept equal to zero?	χ^2_1	2.993
$H_0 : \lambda_0 = 0$	p-val	0.084
Wald test for the restrictions imposed on model III		
Do we lose too much information in model III?	χ^2_3	2.275
$H_0 : \zeta_1 = \zeta_2 = \zeta_3 = 0$	p-val	0.893
Wald test for the restrictions imposed on model V		
Do we lose too much information in model V?	χ^2_6	1.928
$H_0 : k_2 = l_2 = \zeta_2 = n_1 = p_1 = p_2 = 0$	p-val	0.707
Wald tests for listings effects on the unitary risk premium		
Do the three derivatives listings affect the GARCH process only?	χ^2_9	16.009
$H_0 : l_i = n_i = p_i = 0, \forall i$	p-val	0.067
Does the unitary risk premium vary over the sample?	χ^2_{11}	24.864
$H_0 : k_2 = k_3 = l_i = n_i = p_i = 0, \forall i$	p-val	0.009
Are there any derivatives listings effects over the sample?	χ^2_{12}	20.072
$H_0 : \zeta_i = l_i = n_i = p_i = 0, \forall i$	p-val	0.066

This table provides likelihood ratio tests and Wald statistics computed on the most general parameterization (model I) that lead us to keep the intercept in subsequent models and to choose the parsimonious model V. Parameterizations of these nested models are given in Table IV.

[30] Likelihood ratio tests are not adequate if the distribution is not normal. We however report the statistics under χ^2 for information.

Table VI. Quasi-maximum likelihood of the conditional SBC index excess return model with a time-varying unitary market price of risk

	Model I		Model V		Model HP	
λ_0	−0.450 (−0.99)		−0.297 (−1.07)		−3.896 (−1.03)	
a and a^2 respectively	0.278 (3.49)	0.077 (1.75)	0.273 (3.60)	0.075 (1.80)	0.210 (3.00)	0.044 (1.50)
b and b^2 respectively	0.883 (11.92)	0.780 (5.96)	0.871 (14.23)	0.759 (7.12)	0.688 (2.66)	0.473 (1.33)
k_1 (constant)	−2.416 (−1.09)		−2.446 (−2.80)		0.323 (0.34)	
k_2 (Delta Euro)	4.513 (0.35)					
k_3 (LDWR)	−0.242 (−1.34)		−0.362 (−2.01)		0.067 (0.31)	
Individual stocks options listings (19th May 1988)						
ζ_1	−0.351 (−1.40)		−0.424 (−1.84)		−0.639 (−1.22)	
l_1 (constant)	1.787 (0.81)		1.658 (2.29)		0.641 (2.36)	
l_2 (Delta Euro)	−4.264 (−0.33)					
l_3 (LDWR)	0.483 (1.82)		0.661 (2.22)		0.177 (0.70)	
Stock index options listing (7th December 1988)						
ζ_2	0.143 (0.64)					
n_1 (constant)	−0.299 (−0.13)					
n_2 (Delta Euro)	−6.133 (−0.47)		−1.393 (−1.30)		−0.957 (−1.10)	
n_3 (LDWR)	0.602 (2.33)		0.684 (2.79)		0.334 (1.14)	
Stock index futures listing (9th November 1990)						
ζ_3	−0.156 (−1.28)		−0.161 (−1.39)		0.007 (0.08)	
p_1 (constant)	0.797 (0.36)					
p_2 (Delta Euro)	−5.127 (−0.39)					
p_3 (LDWR)	0.392 (1.27)		0.680 (2.10)		0.173 (0.89)	
Log likelihood	−505.85		−507.74		−487.56	

This table gives QML estimates of models I, V and HP. These models capture derivatives listings' effects through dummy variables on both the variance process and the unitary risk premia (a subset of them for models V and HP). The constituents of the information set Z_t for the models I and V are a constant, the first differenced Euro CHF one month rate series (DELTA EURO) and the lagged excess returns of the world stock index over the Euro CHF one month rate series (LWXR). For model HP, the constituents of the information set Z_t are a constant, the Hodrick-Prescott filtered first differenced Euro CHF one month rate series (DELTA EURO) and the Hodrick-Prescott filtered lagged excess returns of the world stock index over the Euro CHF one month rate (LWXR) series. The smoothing factor for both series is equal to 10 and the two filtered series are considered as exogenous. k, l, n and p are hence 3×1 vectors. The variance process is a GARCH (1,1) which enables each derivatives listing to modify the constant of the GARCH process through $\zeta_i . D_i$ is a dummy which equals 1 between the first and the second derivative introductions (i.e. between the 19th May 1988 and the 7th December 1988) and 0 elsewhere. D_2 is a dummy which equals 1 between the second and the third derivative introductions (between the 7th December 1988 and the 9th November 1990). D_3 is a dummy which equals 1 after the third derivative introduction (9th November 1990). Parameters constrained to zero due to the specification of model V are left blank. While we estimated the coefficients a and b, we also report the value for a^2 and b^2 for ease of comparison. Delta method was used to compute the t-stats of these latter coefficients. Robust t-stats are provided in () under the estimates.

$$r_{M,t} = \hat{\lambda}_0 + \exp\left(\left(k' + D_{1,t}.l' + D_{2,t}.n' + D_{3,t}.p'\right).Z_t\right).\hat{\sigma}^2_{M,t} + \hat{\varepsilon}_{M,t}, \tag{A.4}$$

$$\hat{\sigma}^2_{M,t} = \hat{\sigma}^2_0\left(1 - a^2 - b^2\right) + a^2.\hat{\varepsilon}^2_{M,t-1} + b^2.\hat{\sigma}^2_{M,t-1} + \zeta_1.D_{1,t} + \zeta_2.D_{2,t} + \zeta_3.D_{3,t}, \tag{A.5}$$

$$\hat{\varepsilon}_{M,t} = \Im_{t-1} \sim i.d.\left(0; \hat{\sigma}^2_{M,t}\right).$$

Table VII. Average derivatives listings' effects

Panel A	
Mean and variance of the time-varying unitary risk premium	
Over the entire sample[31]	
$\bar{\lambda}_t = \text{mean of} \exp\left[\left(k' + D_{1,t}.l' + D_{2,t}.n' + D_{3,t}.p'\right).Z_t\right]$	0.1447
$\text{Var}(\lambda_t)$	0.1583
Mean of the "base" time-varying unitary risk premium	
"Base" unitary risk premium over the entire sample (derivatives listing effects are not taken into account)	
$\frac{1}{253}\sum_{1}^{253}\exp\left[k'.Z_t\right]$	0.1098
"Base" unitary risk premium after the third introduction (derivatives listing effects are not taken into account)	
$\frac{1}{105}\sum_{149}^{253}\exp\left[k'.Z_t\right]$	0.1066
Mean of the time-varying unitary risk premium after the third introduction	
Unitary risk premium after the third introduction (listing effects are included here)	
$\frac{1}{105}\sum_{149}^{253}\exp\left[\left(k' + p'\right).Z_t\right]$	0.1007
Panel B	
First derivatives average listing effect on the unitary risk premium (individual stock options listing)	
$\frac{1}{28}\sum_{20}^{47} l'.D_{1,t}.Z_{t-1}$	1.8244
note that	
$l'.D_{1,t}.Z_{t-1} = \begin{cases} l_1 + l_3.LWXR_{t-1}, & 20 \leqslant t < 48 \\ 0 \text{ elsewhere} \end{cases}$	
Second derivatives average listing effect on the unitary risk premium (stock index options listing)	
$\frac{1}{101}\sum_{48}^{148} n'.D_{2,t}.Z_{t-1}$	−0.1545
in the same way	
$n'.D_{2,t}.Z_{t-1} = \begin{cases} n_2.Delta\ Euro_{t-1} + n_3.LWXR_{t-1}, & 48 \leqslant t < 149 \\ 0 \text{ elsewhere} \end{cases}$	
Third derivatives average listing effect on the unitary risk premium (stock index futures listing)	
$\frac{1}{105}\sum_{149}^{253} p'.D_{3,t}.Z_{t-1}$	−0.0250
with	
$p'.D_{3,t}.Z_{t-1} = \begin{cases} p_3.LWXR_{t-1}, & 149 \leqslant t \\ 0 \text{ elsewhere} \end{cases}$	

This table provides average derivatives introductions effects on the unitary market risk premium for model V. Parameterizations of this model is given in Table IV. Since the dummies are not overlapping, adding the coefficient of a dummy to its corresponding coefficient in the specification of the unitary risk premium gives the average premium that prevails during the period covered by the dummy. For example, adding the dummy of the third introduction p' to the "base" coefficients k_i gives the coefficient of the unitary risk premium including the listing effects of the three introductions. Panel A provides the implied mean unitary risk premium for different subperiods determined by derivatives introductions while panel B provides the computed averages of the three derivatives listing effects.

[31] If we exclude the period between the first and the second introductions (i.e. excluding observations 20 to 47), this gives $\bar{\lambda}_t = 0.0987$ and $\text{Var}(\lambda_t) = 0.0699$.

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