Can Investment Shocks Explain the Cross-Section of Equity Returns?*

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Abstract

Using two macro-based and one return-based measures of investment-specific technology (IST) shocks, we find that over the 1964–2012 period exposure to IST shocks cannot explain cross-sectional returns spreads based on book-to-market, momentum, asset growth, net share issues, accrual, and price-to-earning ratio. Only one of the two macro-based measures can explain a sizable portion of the value premium over the longer 1930–2012 period. We also find that the IST risk premium estimates are sensitive to the sample period, the data frequency, the test assets, and the econometric model specification. Impulse responses of aggregate investment and consumption indicate potential measurement problems in IST proxies, which may contribute to the sensitivity of IST risk premium estimates and the failure of IST shocks to explain cross-sectional returns.

JEL Classification Codes: G12; O30

Keywords: Investment shocks, cross-sectional returns, value premium

1 Introduction

Investment-specific technology shocks (IST shocks hereafter)—i.e., technological innovations that materialize through the creation of new capital stock—have long been recognized by economists as important determinants of economic growth and business cycle fluctuations.[1] More recently, financial economists have relied on IST shocks as an economically motivated risk factor for explaining properties of asset prices in both the cross section and time series. Theoretically, the effect of IST shocks on asset prices depends crucially on key channels such as investors' preferences toward the resolution of uncertainty (e.g., Papanikolaou (2011)), firms' flexibility in varying capital utilization and their degree of market power (e.g., Garlappi and Song (2016a)). Empirically, the evidence from existing studies that rely on IST shocks to analyze cross-sectional equity returns is mixed. For example, Kogan and Papanikolaou (2014) argue that a *negative* price of risk is needed to explain the value premium—i.e., the fact that high book-to-market (B/M) firms earn higher returns than low B/M firms (see Fama and French (1992)). In contrast, Li (2013) argues that a *positive* price of risk is needed to explain the profitability of momentum strategies—i.e., the fact that stocks with high past returns outperform stocks with low past returns (see Jegadeesh and Titman (1993)). To establish the relevance of IST shocks as an economically motivated risk factor and differentiate among alternative theories, it is therefore important to analyze in depth the empirical evidence on the ability of these shocks to explain cross-sectional returns.

In this paper, we assess the role of IST shocks for cross-sectional asset prices by focusing first on the same return patterns that brought forth the aforementioned disagreement: the value premium and the momentum effect. We then broaden the scope of our empirical analysis by studying the effect of IST shocks on alternative cross-sections represented by portfolios sorted by the following firm characteristics: (i) asset growth rate, (ii) net share issues, (iii) earnings-to-price (E/P) ratio, and (iv) accrual. We conduct our study on value and momentum over a long sample period, from 1930 to 2012, and on the alternative

[1] See, for example Solow (1960), Greenwood, Hercowitz, and Krusell (1997, 2000), Christiano and Fisher (2003), Fisher (2006), and Justiniano, Primiceri, and Tambalotti (2010, 2011).

cross-sections over the more recent sample, from 1964 to 2012, due to data availability. Our empirical analysis adopts measures of investment-specific technology shocks that have been widely used in the macro-finance literature. The first measure, *Ishock*, proposed by Greenwood, Hercowitz, and Krusell (1997), is based on the (quality-adjusted) price of capital goods relative to that of consumption goods and aims to capture shocks to the cost of investment in new capital. The second measure, *IMC*, proposed by Papanikolaou (2011), is based on the stock return spread between aggregate investment and consumption good producers. The third measure, *gIMC*, first used by Kogan and Papanikolaou (2014), is the growth rate spread between aggregate investment and consumption.

Using a standard Fama and MacBeth (1973) two-stage estimation procedure and a broad cross-section of 40 test assets (10 size, 10 B/M, 10 momentum, and 10 industry portfolios) from 1930 to 2012, we obtain positive and significant estimates of the IST risk premium. Combining these findings with the estimates of IST loadings, we infer that only the exposure to *Ishock* can explain a sizable part (up to 62%) of the value premium, while the explanatory power of the other two measures is much weaker (at most 35%). For momentum, we find that the two macro-based IST measures (*Ishock* and *gIMC*) can explain a sizable fraction of momentum profits (up to 46%). However, for all three IST measures, we strongly reject the hypothesis that exposure to IST shock can explain the full magnitude of momentum profits. We confirm that these results are qualitatively similar if we exclude the Great Depression period and limit the analysis to the post–World War II sample from 1948 to 2012.

Our finding that one of the IST measures, *Ishock*, can explain a large fraction of the value premium in the 1930–2012 sample, is broadly consistent with that of Papanikolaou (2011) and Kogan and Papanikolaou (2014), who use IST risk exposure to explain the value effect. However, our analysis also shows that the ability of IST shocks to account for the value effect crucially depends on the sample period used. For example, none of the three IST measures can generate sizable value premia in the post-1963 sample (exposure to IST shocks can explain at most 24% of the value premium, in the annual sample, and at most 4%, in the quarterly sample). Similarly, our finding that macro-based IST measures can explain a sizable fraction of momentum profit in the 1930–2012 sample is broadly consistent with Li (2013), who uses IST shocks to explain momentum profits. However, as for the case of the value premium, the effect of IST shocks on momentum is also sample-dependent. For example, none of the three IST proxies can generate sizable momentum profits in the quarterly post-1963 sample. Exposures to IST shocks can explain at most 2% of momentum profits.

For the more recent 1964–2012 sample, risk exposures to IST shocks fail to explain not only value premium and momentum, but also return spreads of cross sectional portfolios based on: (i) Asset growth rate, (ii) Net share issues, (iii) E/P ratio, and (iv) Accrual. Therefore, while IST shocks appear to have some explanatory power for the value premium in the full sample, their explanatory power for cross sectional equity returns diminishes in the more recent sample. Our results are robust to the use of different test assets in the estimation of the IST price of risk.

Our positive estimates of the IST risk premium from 1930–2012 stand in contrast to the negative estimates that Papanikolaou (2011) and Kogan and Papanikolaou (2013, 2014) obtain using post-1963 data. To understand and reconcile this difference, we first replicate and confirm their negative estimates using post-1963 data and a cross-section of ten book-to-market portfolios. We then document that the inference based on the three proxies of IST shocks depends crucially on both the sample period and the test assets employed. For example, using the ten book-to-market portfolios as the only test assets, the estimated IST risk premium is negative for post-1963 but *positive* for pre-1963 data. Moreover, even for post-1963 data, the IST risk premium is positive when estimated from a cross-section of test assets that is broader than the ten book-to-market portfolios. We further show that the inference on IST risk premium also depends on the econometric model specification. For example, using the post-1963 sample of ten IMC-beta sorted portfolios as test assets, we find that, consistent with Kogan and Papanikolaou (2014), IST risk premium estimates are negative and significant when we ignore the intercept in cross-sectional regressions. However, if we allow for an intercept, the IST risk premium estimates become indistinguishable from zero and the intercept estimate is significant.

For robustness, we finally show that IST estimates from Fama-MacBeth regressions are equivalent to those obtained from the Generalized Method of Moments (GMM) approach.

Our finding that the sign of the IST risk premium differs across test assets and sample periods has important implications for theoretical models. For example, in the general equilibrium model of Papanikolaou (2011), IST risk premia are negative because a positive IST shock induces a drop in consumption, hence increasing marginal utility. This makes an asset with positive IST exposure a "hedge" against consumption risk. On the other hand, Garlappi and Song (2016a) show that if firms can increase their capital utilization upon a positive IST shock, consumption may increase rather than decrease. This in turn implies that marginal utility may be lower upon a positive IST shock. In this case, an asset with positive IST exposure is risky and therefore demands a positive risk premium.[2]

Finally, we document evidence indicating the existence of measurement problems in commonly used IST proxies. For example, we find that while the two macro-based proxies of IST shocks, *Ishock* and *gIMC*, seem to have strong comovement with both the aggregate consumption and investment, such a comovement weakens in the more recent sample. In contrast, the return-based IST proxy, *IMC*, does not comove with either investment or consumption, indicating potentially larger measurement errors for this IST proxy. These findings call for more effort in addressing the measurement problems in IST proxies.

Our paper is closely related to the recent finance literature that studies the effect of IST shocks on asset prices. Papanikolaou (2011) is the first to study the implications of these shocks for asset prices in the cross-section of stocks. He introduces IST shocks in a two-sector general equilibrium model and shows how financial data can be used to measure IST shocks at a higher frequency. In a partial equilibrium setting, Kogan and Papanikolaou (2013, 2014) explore how IST shocks can explain the value premium as well as other return patterns in the cross-section that are associated with firm characteristics, such as Tobin's Q, past investment, earnings-price ratios, market betas, and idiosyncratic

[2]Another important channel that affects the sign of the IST risk premium is the investors' preferences toward the resolution of uncertainty. For example, under the preference specification adopted by the long-run risk literature (e.g., Bansal and Yaron (2004), Ai, Croce, and Li (2013), Croce (2014)), Garlappi and Song (2016a) show that IST shocks demand a positive risk premium. This contrasts with the negative IST risk premium under the preference specification in Papanikolaou (2011).

volatility of stock returns. Li (2013) proposes a rational explanation of the momentum effect in the cross-section by using investment shocks as the key risk factor. Yang (2013) uses investment shocks to explain the commodity basis spread.[3]

We make three contributions to the literature on cross-sectional asset pricing. First, we provide a thorough empirical analysis of the effect of investment-specific shocks on the value premium, momentum profits and other significant cross-sectional return patterns. The long sample period (1930–2012) we consider in this paper offers an "out-of-sample" analysis that complements existing studies in which the focus is mainly on relatively recent data (post-1963). Second, the new evidence we provide sheds some light on the economic mechanisms proposed in existing general equilibrium models with IST shocks, and therefore enhances our understanding of the effect of IST shocks on asset prices. Third, we highlight the existence of potentially severe measurement problems in commonly used IST proxies and call for more effort in future research that aims to use IST shocks to study the behavior of asset prices and macroeconomic quantities.

The rest of the paper is organized as follows. Section 2 describes the data. Section 3 provides empirical evidence from cross-sectional analysis. Section 4 compares our empirical findings on the IST risk premium with the existing literature. Section 5 provides further discussions on the measurement issues related to IST proxies and future research directions. Section 6 concludes. Appendix A contains details of the data we use.

2 Data

In this section, we briefly describe the construction of the empirical measures of IST shocks and report their statistical properties. Due to data availability, we consider two samples in our analysis. The first sample consists annual data from 1930 until 2012, for which both B/M and momentum portfolios are available. To allay the concern that this

[3]Our paper is also broadly related to a large body of literature that uses heterogeneity in firms' investment decisions and their exposures to disembodied productivity shocks to explain cross-sectional returns, as pioneered by Cochrane (1996) and Berk, Green, and Naik (1999). Significant contributions to this literature that are closely related to the cross-sections we study include Carlson, Fisher, and Giammarino (2004) and Zhang (2005) for the value effect, and Sagi and Seasholes (2007) and Liu and Zhang (2008, 2014) for the momentum effect.

sample period contains the tumultuous time of the Great Depression, for robustness, we also consider post–World War II data, both at the annual and quarterly frequency. The second sample consists data from 1964 to 2012, for which all the cross-sectional returns we consider are available at both annual and quarterly frequency. Appendix A contains a more detailed description of all the data we use.

2.1 Measures of IST shocks

Because IST shocks are not observable, we need to rely on plausible empirical proxies. We choose three IST proxies that are designed to capture different aspects of IST shocks. The first proxy, *Ishock*, focuses on the effect of IST shocks on the price of capital goods. The second proxy, *IMC*, focuses on the effect of IST shocks on stock returns of firms in the investment vs. consumption sector. The third proxy, *gIMC*, focuses on the effect of IST shocks on the growth of aggregate investment and consumption.[4] By relying on a variety of proxies we aim to provide a comprehensive and thorough investigation on the effect of IST shocks on the cross section of equity returns. We now describe in detail the construction of these three proxies.

2.1.1 IST proxy based on the relative price of capital goods: ***Ishock***

Our first measure of IST shocks, *Ishock*, was originally proposed by Greenwood, Hercowitz, and Krusell (1997) and is the change in the price of investment goods relative to that of nondurable consumption goods. Specifically, for period t, *Ishock* is defined as

$$Ishock_t = -\left(\ln\left(P_I/P_C\right)_t - \ln\left(P_I/P_C\right)_{t-1}\right), \tag{1}$$

where P_I is the price deflator for equipment and software of gross private domestic investment, and P_C is the price deflator for nondurable consumption goods. The price deflator for nondurable consumption goods, P_C, is from the National Income and Prod-

[4]We also repeat the analysis of this paper using a fourth IST proxy that we construct from the first principal component extracted from these three proxies. The results are qualitatively similar to those inferred from the *Ishock* and *gIMC* proxies and are available upon request.

uct Accounts (NIPA) tables. The price deflator for investment goods, P_I, is from the quality-adjusted series of Israelsen (2010).[5]

The idea behind this measure is intuitive. If a new investment-specific technology improves the production of investment goods, the increased supply of investment goods would lead to a drop in the price of investment goods relative to consumption goods. That is, a positive IST shock leads to a reduction in the relative price of equipment, and therefore, a positive value for *Ishock*.

2.1.2 IST proxy based on return spreads: *IMC*

Our second measure of IST shocks, *IMC*, was originally proposed by Papanikolaou (2011) and is the return difference between investment and consumption sectors,

$$IMC_t = r_t^I - r_t^C, \tag{2}$$

where r_t^I and r_t^C are the returns of firms producing investment goods and consumption goods, respectively. The classification of a firm as belonging to the consumption or investment sector is based on the procedure of Gomes, Kogan, and Yogo (2009), who classify each Standard Industry Classification (SIC) code into either investment or consumption sector based on the 1987 benchmark input-output accounts.

The rationale for using the *IMC* return as a measure of IST shocks is that, under the assumptions of the two-sector general equilibrium model of Papanikolaou (2011) or the vintage capital partial equilibrium model of Kogan and Papanikolaou (2013, 2014), firms producing investment goods (investment firms) and consumption goods (consumption firms) have the same loadings on the neutral productivity shock, but different loadings on IST shocks. If so, the return spread between investment and consumption firms loads only on IST shocks and can therefore be used as an alternative proxy for these shocks. Because it is constructed from financial markets data, the *IMC* measure has the advantage of being available at a higher frequency than *Ishock*.

[5] We are grateful to Ryan Israelsen for sharing with us the annual series of quality-adjusted prices from 1930 to 2012.

2.1.3 IST proxy based on sectoral growth spreads: *gIMC*

Our third measure of IST shocks, *gIMC*, is the growth rate difference between aggregate investment and consumption,

$$gIMC_t = g_t^I - g_t^C, \tag{3}$$

where g_t^I and g_t^C are the log growth rates of aggregate investment and consumption, respectively. The intuition behind the *gIMC* measure is similar to that of *IMC*. In a model with both neutral TFP shocks, affecting equally both investment and consumption, and capital-embodied IST shocks, affecting only investment, the growth difference between investment and consumption should be closely related to IST shocks. We take the spread, *gIMC*, between growth rates as a proxy for IST shocks. Note that *gIMC* is equivalent to the growth rate in the investment-to-consumption ratio used by Kogan and Papanikolaou (2014).

2.2 Time-series properties of IST measures

Table 1 reports summary statistics of the three IST measures and their correlations with macro factors (the growth rates of consumption expenditures, *GDP*, and *TFP*) and return factors (market, size, value, and momentum). The two macro-based measures, *Ishock* and *gIMC*, are available from 1930 at the annual frequency and from 1948 at the quarterly frequency. We also reports results for the post-1963 period at both annual and quarterly frequency.

The annual mean and standard deviation for *Ishock* over the entire sample period (1930–2012) are 3.45% and 3.68%, respectively. The average *Ishock* is positive and significant for all the sample periods we report. In other words, according to the *Ishock* measure, we do observe improvement in investment-specific technology in the US economy. In contrast, the averages of *IMC* and *gIMC* are not significant across different sample periods. This lack of significance is potentially due to the fact the volatility of these measures is much higher than that of *Ishock*.

The contemporaneous correlations of *Ishock* with the growth rate of personal consumption expenditure (*PCE*), *GDP*, and *TFP* are, respectively, 0.20, 0.46, and 0.22 over the entire 1930–2012 annual sample and become statistically insignificant in the postwar 1948–2012 sample and over the more recent 1964–2012 sample. Unlike *Ishock*, *IMC* does not exhibit any significant correlation with the macro factors in the annual time series. However, in the quarterly time series spanning the 1948–2012 period, *IMC* is positively correlated with all three macro factors. The growth spread, *gIMC*, is positively correlated with the three macro factors across different sample periods.

The correlation between IST proxies and the return factors (market (MKT), size (*SM-B*), value (*HML*), and momentum (*UMD*)) are time-varying. For example, the correlations of *Ishock* with Fama-French 3-factors are all positive and significant for the 1930–1963 period, and they all turn negative over the 1964–2012 period. A similar switch also happens for the correlation between *IMC* and *HML*.[6] Note also that the two macro-based proxies, *Ishock* and *gIMC*, are positively correlated with the momentum factor, however, the financial-based proxy, *IMC* is uncorrelated with momentum.

The last three columns of Table 1 report the correlation matrix for the three IST measures. The highest correlation is 0.34 between *Ishock* and *gIMC* over the 1964–2012 annual sample period. Overall, the low level of correlation among IST proxies indicates the existence of a potential measurement issue with these proxies, as we discuss in Section 5.

In summary, the analysis in this section shows that the measures of IST shocks introduced in Subsection 2.1 are pro-cyclical and exhibit positive correlation with return factors. The subsample analysis suggests that the statistical properties of these measures are time-varying. We formally investigate the asset pricing implications of this time variation in Section 3.

[6]The sign change in the correlation between *HML* and the two IST proxies is interesting. Together with the findings in the prior literature documenting that the CAPM holds well in the early subsample but not in the late one (see, e.g., Davis, Fama, and French (2000), Campbell and Vuolteenaho (2004), Ang and Chen (2007), and Fama and French (2006)), the evidence in Table 1 seems to indicate that *HML* experiences some kind of "structural break" around 1963. One plausible explanation for such a change in *HML* is the "changing nature" of book-to-market portfolios. For example, Chen (2014) finds a similar structural break in the relative growth rate of cash-flow of value vs. growth firms: dividends of value stocks grow faster (slower) than those of growth stocks in the pre-1963 (post-1963) sample.

2.3 Cross-sectional test assets

To estimate the risk premium of IST shocks, we choose a cross-section of 40 test portfolios: size deciles, book-to-market deciles, momentum deciles, and ten industry portfolios. The first 30 portfolios have been used in the literature as test assets for the estimation of risk premia of aggregate risk factors (see for example, Liu and Zhang (2008) and Cooper and Priestley (2011)). Because the impact of investment shocks is likely to differ across industries, we also include ten industry portfolios in the set of test assets. These 40 test portfolios are all available for the 1930–2012 sample period. To assess the robustness of the IST risk premium estimates, we also consider an alternative set of 40 test portfolios that include: profitability deciles, asset growth deciles, volatility deciles, and net share issues deciles. These alternative test portfolios are available only for the more recent 1964–2012 sample period.

Due to data availability, we study the effect of IST exposure on cross-sectional return spreads generated by book-to-market (high minus low B/M) and momentum (winners minus losers) in the 1930–2012 sample. For the more recent 1964–2012 sample period, we broaden our analysis to include four additional return spreads that are generated by: (i) asset growth (low minus high growth), (ii) net share issues (low minus high issues), (iii) earning-to-price (high minus low E/P), and (iv) accrual (low minus high accrual). We study the explanatory power of the IST shocks on these six cross-sections, which all show significant return spreads in the sample.

3 IST shocks and cross sectional equity returns

In this section, we empirically investigate whether well-known cross-sectional return patterns in equity returns can be linked to firms' exposure to IST shocks. We first estimate the risk premium of IST shocks via Fama-MacBeth regressions in Section 3.1. We then assess whether exposure to IST shocks can explain the observed cross sectional variation in equity returns over two sample periods. In Subsection 3.2, we study the value premium and momentum spread for the 1930–2012 annual sample. In Subsection 3.3, we expand our analysis to all the six cross-sections described in Subsection 2.3 for the more recent 1964–2012 sample period. In Section 3.4, we provide further robustness analysis by using alternative test assets in estimating the IST risk premium and the post-WWII sample for value and momentum, where both annual and quarterly data are available.

3.1 The IST risk premium

To estimate the IST risk premium, we rely on standard two-stage Fama and MacBeth (1973) regressions in which we use as proxies for IST shocks the measures described in Section 2.1. We estimate the IST risk premium for two samples. The first is the full 1930–2012 sample, for which we have cross-sectional return data on the B/M and momentum portfolios that we study in Section 3.2. The second is the more recent 1964–2012 sample, for which we have return data for all the six cross-sections that we study in Section 3.3. In both cases, we use the respective samples (i.e., either 1930–2012 or 1964–2012) in the first-stage time-series regressions to estimate the risk loadings of the 40 test assets described in Section 2.3. We then use the same test assets in the second-stage cross-sectional regressions to estimate the IST risk premium.

When estimating the IST risk premium, we control for a common, disembodied, aggregate factor in the economy which we measure using three different proxies: (i) the market excess return (MKT), (ii) the growth rate of *TFP*, and (iii) the growth rate of consumption (gC). For each proxy of IST shocks, we estimate the IST risk premium both in univariate and bivariate models, where we control for the common aggregate factor. Details of the construction of MKT, *TFP*, and gC are in Appendix A. Since the betas used in the second-stage regressions are estimated, we correct the standard errors and t-statistics following Shanken (1992). In addition, we also adjust the t-statistics in the second-stage estimation for potential autocorrelation and heteroskedasticity following Newey and West (1987).

Panel A of Table 2 reports the risk premium estimates from the second stage of Fama-MacBeth regressions using the full 1930–2012 annual sample. The risk premium estimates from *Ishock* are positive and significant. For example, in the univariate model

(model (1a)), the risk premium for *Ishock*, λ_{Ishock}, is 3.77% per year with a t-statistic of 2.35. The results are similar after we control for one of the three common factors in bivariate cross-sectional regressions (models (1b), (1c), and (1d)).

The risk premium estimates from *IMC* (λ_{IMC}) are also positive in both univariate and bivariate models. The point estimates vary from 0.55% (model (2c)) to 3.83% (model (2a)). However, none of the estimates is statistically significant. The risk premium estimates from *gIMC* (λ_{gIMC}) are also positive in both univariate and bivariate models. However, the point estimates and the statistical significance are model-dependent. For example, estimates of λ_{gIMC} are high and significant after controlling for MKT (9.59% in model (3b) with t-stat of 3.05) or *TFP* (8.28% in model (3c) with t-stat of 2.38). In contrast, in both the univariate model (3a) and the bivariate model (3d) with gC as the second factor, the λ_{gIMC} estimates are relatively low and insignificant.

Panel B of Table 2 reports the risk premium estimates from the more recent 1964–2012 annual sample. In contrast to the full 1930–2012 sample, the IST risk premium estimates are mostly insignificant. Only in the bivariate models of *gIMC* with MKT (model (3b)) and *TFP* (model (3c)), the IST risk premium estimates are positive and statistically significant.

In summary, the results in Table 2 indicate that, based on annual data from 1930 to 2012, IST shocks demand a positive risk premium. However, the statistical significance of the estimates depends on the empirical proxy for IST shocks (*Ishock*, *gIMC*, or *IMC*), the regression model considered, and the sample period.

3.2 Cross-sectional returns in the 1930–2012 period

In this subsection, we study the effect of IST shocks on cross-sectional returns in the 1930–2012 sample. Data availability allows us to focus only on two cross-sections, namely, B/M and momentum portfolios. We expand our analysis to broader cross-sections in the more recent sample period in the next subsection.

The contribution of IST shocks to an asset's risk premium is the product, $\lambda_{IST} \times \beta_{IST}$, of the IST risk premium, λ_{IST}, and the IST loading of the asset's returns, β_{IST}. In the previous subsection we determined the IST risk-premium λ_{IST} via a two-stage Fama-MacBeth regression. To assess whether IST shocks can explain the value and momentum effect, it is necessary to estimate the IST loadings β_{IST} of book-to-market and momentum portfolios. In this section, we compute the loadings via time series regressions over the entire 1930–2012 sample period.

Panel A of Table 3 reports the returns and IST loadings of ten book-to-market sorted portfolios for the 1930–2012 annual sample. The portfolio excess returns r increase from 6.61%, for the growth portfolio (low decile), to 13.67%, for the value portfolio (high decile), implying a statistically significant difference of 7.05% per annum (with t-stat=2.56). The IST beta loadings are obtained from time series regressions of portfolio excess returns on the chosen measure of IST shock, i.e., *Ishock*, *IMC*, or *gIMC*. In general, IST loadings of value stocks are higher than those of growth stocks. However, the difference between IST betas of value and growth portfolios are not statistically significant. For example, the univariate beta loadings on *Ishock*, β_{Ishock}, increase from 0.22, for the growth portfolio, to 1.19, for the value portfolio. Similarly, the univariate beta loadings on *IMC*, β_{IMC}, increase from 0.66, for the growth portfolio, to 1.09, for the value portfolio. The general pattern is similar for bivariate betas (not tabulated), obtained from time-series bivariate regressions that include, in addition to the IST proxy, either the market factor, *MKT*, the growth rate of total-factor productivity, *TFP*, or consumption growth, *gC*.

Panel B reports the corresponding quantities for the momentum deciles. The excess returns, r, increase from 0.8%, for the portfolio of losers, to 15.70%, for the portfolio of winners, implying a statistically significant difference of 14.9% per annum (with t-stat=5.47). Betas on *Ishock* and *gIMC* also show an increasing pattern from losers to winners. For example, the univariate beta loadings on *Ishock*, β_{Ishock}, increase from -0.47, for the loser portfolio, to 1.22 for the winner portfolio, resulting in a beta difference of 1.69 with t-stat of 1.80. The significance of the beta difference between winners and losers is much higher for *gIMC* (t-stat=2.56). In contrast, *IMC* betas are *lower* for winner portfolios than for loser portfolios. For example, the univariate beta loadings on *IMC*, β_{IMC}, decrease from 1.13, for the loser portfolio, to 0.83 for the winner portfolio, resulting in an insignificant beta for winners-minus-losers. As for book-to-market portfolios, the bivariate IST betas for momentum portfolios (not tabulated) show a similar pattern as the univariate betas. In summary, for all IST measures, with the exception of *IMC*, IST betas of value portfolios (winners) are larger than those of growth portfolios (losers).

Note that the IST betas reported in Table 3 are statistically insignificant when the two macro-based measures (*Ishock* and *gIMC*) are used, but highly significant if the return-based measure (*IMC*) is used. Our further analysis indicates that the insignificance of the time series IST beta estimation is a result of time variation in the IST betas. Specifically, the *Ishock* betas are positive and significant in the pre-1963 subsample but negative and significant in the post-1963 subsample. The full sample beta is effectively the average of the betas over the two sample periods. The opposite signs of betas over the two sample periods explain the low statistical significance of the full sample estimates. This time variation in IST betas indicates the importance to investigate the IST pricing effect in subsamples, as we do in Subsection 3.3. In addition, the IST betas reported in Table 3 are not monotonic in the sorting characteristics, indicating potential measurement problems in IST proxies, which we discuss further in Section 5.

The above estimates of IST betas, together with the IST risk premium estimates of Section 3.1, allow us to calculate the fraction of value premium and momentum profits that can be explained by exposure to IST shocks. The *realized* value premium, HML, is the difference in the return between the High and Low book-to-market portfolios. Over the 1930–2012 sample period the value premium is 7.05% per annum as reported in Table 3. The component of value premium explained by IST shocks, which we denote by $\beta_{IST}\lambda_{IST}$, is equal to the product of: (i) the spread in betas β_{IST} between value and growth portfolios (from Table 3), and (ii) the estimate of the risk premium λ_{IST} for IST shocks (from Panel A of Table 2). We refer to the quantity $\beta_{IST}\lambda_{IST}$ as the *expected value premium* from exposure to IST shocks. For example, using univariate *Ishock* betas, β_{Ishock}, from Panel A of Table 3, and the IST risk premium λ_{Ishock} from the single factor model (1a) in Panel A of Table 2, the IST component of the value premium is $\beta_{IST}\lambda_{IST} = 0.98 \times 3.77\% = 3.69\%$ per annum. Given an observed value premium of 7.05%, this means that *Ishock* can explain 52% of the observed value premium. We follow a similar procedure to determine the contribution of IST shocks to momentum profits. For the cases with two factors, the model-implied expected return $\widehat{HML}$ (or $\widehat{WML}$) includes the contributions from both the IST risk and the second risk factor (*MKT*, *TFP* or *gC*) calculated in a similar fashion.

Panel A of Table 4 reports the results for the value premium. Three points are worth mentioning. First, *Ishock* risk exposure explains 52% of the value premium (column labeled $\frac{\beta_{IST}\lambda_{IST}}{HML}$) in the univariate model, close to 40% in bivariate models that use *MKT* and *TPF* as aggregate factors, and 62% in the bivariate model that uses *gC* as the aggregate factor. The column labeled "t(diff1)" reports the t-statistics for the test of the null hypothesis that IST shocks explain the value premium ($\beta_{IST}\lambda_{IST} - HML = 0$).[7] The t-statistics for these tests reveal that, with a 5% significance level, we cannot reject the hypothesis that exposure to *Ishock* explains the value premium. Second, exposure to *IMC* can only explain a small fraction of the value premium (ranging from 3% to 23%, depending on models), and we reject the hypothesis that *IMC* risk exposure can explain the value premium. Finally, exposures to *gIMC* generate *negative* value premium and therefore fail to explain the observed positive value premium.

Panel B reports the corresponding results for momentum profits. Exposures to *Ishock* explain about 28% to 43% of momentum profits, but we reject the hypothesis that *Ishock* exposures can explain the magnitude of momentum profits (see the t-statistics in column "t(diff1)"). From the *IMC* proxy of IST shocks, we typically infer *negative* expected momentum profits ($\beta_{IST}\lambda_{IST}$). Finally, *gIMC* generates positive expected momentum profits, but we reject the hypothesis that it can explain the magnitude of these profits.

Our finding that, using annual data, *Ishock* exposures can explain a sizable fraction (ranging from 28% to 43%) of momentum profits is broadly consistent with Li (2013), who claims that IST shocks can explain the momentum effect. However, as discussed above,

[7] Similarly, the column labeled "t(diff2)" reports the t-statistics for the test of the null hypothesis that the bivariate models explain the value premium ($\widehat{HML} - HML = 0$).

we reject the hypothesis that exposure to *Ishock* can explain the magnitude of momentum profits. Moreover, our analysis provides two new findings that challenge the claim that IST exposures explain momentum. First, in contrast to *Ishock*, the *IMC* measure does not have any explanatory power for momentum. Second, as we will show in Subsection 3.4, the explanatory power of *Ishock* is very low when using post–WWII data, explaining at most 15%, when using annual data, and only 4%, when using quarterly data.

In summary, our analysis suggests that, over the entire 1930–2012 sample, *Ishock* exposures can explain a large fraction of the value premium, *IMC* exposures can only explain a much smaller fraction, and *gIMC* exposures generate a counterfactual *growth* premium. Finally, none of the three IST exposures can capture the magnitude of momentum profits.

3.3 Cross-sectional returns in the more recent 1964–2012 period

To broaden the scope of our study, we extend the analysis of the previous subsection to four additional cross sections of assets sorted by: (i) asset growth rate, (ii) net share issues, (iii) E/P ratio, and (iv) accrual. We chose these portfolios because they exhibit significant cross sectional return spreads.

For all these cross sections, Table 5 follows the same structure of Table 4 and reports the expected return spread attributable to exposures to IST risks. In estimating the expected returns, the risk premium for IST shocks, λ_{IST}, is taken from Panel B of Table 2. Since the results are similar across univariate and bivariate models, we report only the univariate results for simplicity.

The results reported in Table 5 indicate that, for the recent 1964–2012 sample, risk exposures to IST shocks fail to explain not only the B/M and momentum effects, but also the other four cross-sectional return spreads. We conclude that, even though IST shocks have some explanatory power for the value premium in the full sample, their explanatory power for cross sectional equity returns dwindles in the more recent sample.

3.4 Robustness

In this subsection we assess the robustness of our results along two dimensions. First, to allay the concern that the risk premia estimates from the 1930–2012 sample are affected by the episodes of high volatility of the Great Depression, we repeat the analysis of Subsection 3.2 on post–World War II data at both the annual and quarterly frequency (1948Q1–2012Q4). Second, we assess the robustness of our results to the choice of test assets used in the estimation of the IST price of risk. In particular, we repeat the analysis by choosing the alternative set of 40 test assets described in Subsection 2.3, which includes decile portfolios sorted along four different characteristics: (i) profitability, (ii) asset growth rate, (iii) volatility, and (iv) net share issues. Because most of these portfolios require accounting data, our robustness analysis using alternative test assets is limited to the post-1963 sub-sample. To save space, we summarize below the main findings from our robustness analysis.

The IST risk premium estimates from the 1948–2012 period are qualitatively similar to those from the 1930–2012 period. That is, these estimates tend to be positive in both univariate and bivariate models and across all three proxies considered. However, the statistical significance is generally lower than those reported in Table 2. Using these IST risk premia estimates we find that, consistent with our original analysis, exposure to IST shocks do not explain cross sectional return spreads in the post-WWII subsample.

The IST risk premium estimates based on the alternative set of 40 portfolios for the 1964–2012 sample are reported in Table 6. The differences between Panel A of Table 6 and Panel B of Table 2 reflect the sensitivity of IST risk premium estimates to the test assets. Comparing the two panels, the estimates for *Ishock* and *IMC* are qualitatively similar, but the risk premium on *gIMC* changes from positive and significant, in Table 2, to negative and insignificant, in Table 6. Panel B of Table 6 reports the risk premium estimates using quarterly data from the 1964Q1–2012Q4 sample. In contrast to the annual data in Panel A, the quarterly *Ishock* risk premium estimates are positive and highly significant. Using these alternative test assets to estimate the IST risk premium, we confirm the findings in Subsection 3.3, namely, that risk exposures to IST shocks fail to explain a large fraction of all the six cross-sectional return spreads considered.

In summary, excluding the Great Depression from our analysis does not change in a significant way our main conclusion from the analysis based on the annual 1930–2012 sample in Subsection 3.2, and using alternative test assets for 1964–2012 sample provides qualitatively similar results as those reported in Subsection 3.3.

4 Comparison with the existing literature

As we discussed in Subsection 3.2, our finding that macro-based IST measures (e.g., *Ishock* and *gIMC*) can explain a sizable fraction of momentum profit in the 1930–2012 annual sample is broadly consistent with Li (2013), who argues that *Ishock* can explain momentum profits. However, our analysis also shows that the effect of IST shocks on momentum is sample-dependent.

Our finding that one of the IST measures, *Ishock*, can explain a large fraction of the value premium in the 1930–2012 annual sample, is broadly consistent with that of Papanikolaou (2011) and Kogan and Papanikolaou (2014), who use IST risk exposure to explain the value effect. However, similar to the case of momentum, our analysis also shows that the ability of IST shocks to account for the value effect depends on the sample period and on the IST proxy used. Setting aside these sensitivity issues, our findings differ from the existing studies in a very important *qualitative* dimension. While our estimate of the IST risk premium is *positive* (see Table 2), Papanikolaou (2011) and Kogan and Papanikolaou (2013, 2014) document a *negative* risk premium of IST shocks.

This difference in the sign of the IST risk premium has at least two important implications. First, a positive IST risk premium implies that in a representative-agent general equilibrium model, the agent's marginal utility is low under a positive IST shock and therefore, assets whose payoff positively correlate with IST shocks are risky. On the other hand, a negative IST risk premium implies that positive IST beta assets are a "hedge" against consumption risk. Second, the sign of the cross-sectional risk premium due to IST risk exposure depends directly on the sign of the IST risk premium. To understand the source of this discrepancy with regard to the sign of the IST risk premium estimates, it is therefore important to compare our findings with those reported in the existing literature.

There are three main differences between our analysis and that of Papanikolaou (2011) and Kogan and Papanikolaou (2013, 2014). First, their analysis is based on post-1963 annual data, while we rely on a longer annual sample spanning from 1930 to 2012, and, in the post-1963 period, consider also quarterly data. Second, their estimate of the IST risk premium is based on cross-sections of test portfolios that are different from the set of 40 test portfolios that we use.[8] Third, their estimates of the IST risk premium are obtained via the Generalized Method of Moments (GMM) while we use a two-stage Fama-MacBeth methodology.

In Subsection 3.1 we have shown that the IST risk premium estimates are mostly positive even in the post-1963 sample (see Panel B of Table 2). This indicates that, in order to understand our findings in light of the existing literature, besides the sample period difference, it is important to investigate the role played by the choice of test assets and the econometric methodology. We undertake such a task in this section. In Subsection 4.1, we show that the IST risk premium estimates are sensitive to the test assets used. In particular, we find that if, instead of the 40 portfolios used in Section 3, we restrict the set of test assets to only ten book-to-market portfolios, the IST risk premium estimates can switch from positive, in the early sample, to negative, in the more recent sample. Similarly, we show that the IST risk premium estimates may be sensitive to the econometric model specification. For example, using the post-1963 sample of IMC-beta sorted portfolios, the IST risk premium estimates can switch from negative and significant, under the assumption of a zero cross-sectional intercept, to indistinguishable from zero if we allow nonzero intercept in the cross-sectional estimation. Finally, in Subsection 4.2,

[8]Specifically, Kogan and Papanikolaou (2013) estimate the IST risk premium using a cross-section of 20 portfolios obtained by taking the first, second, ninth, and tenth decile portfolios from each of the following five cross-sectional sorts: Tobin's q, investment-to-capital ratio (I/K), price/earning ratio (P/E), market beta ($MBETA$), and idiosyncratic volatility ($IVOL$). Kogan and Papanikolaou (2014) estimate the IST risk premium using three separate cross-sections: (i) ten IMC beta sorted portfolios, (ii) ten book-to-market sorted portfolios, and (iii) 30 Fama and French (1997) industry portfolios.

we compare our Fama-MacBeth estimates with those from GMM and show that the two approaches generate equivalent point estimates and differ only slightly in their statistical significance.

4.1 Different test assets in risk premium estimation

4.1.1 Book-to-market portfolios only

Panel A of Table 7 reports the IST risk premium estimates obtained from Fama-MacBeth regressions on ten book-to-market portfolios over the 1964–2012 annual subsample. This set of test assets and sample period have been used by Kogan and Papanikolaou (2014) in their estimation of the IST risk premium. Consistent with Papanikolaou (2011) and Kogan and Papanikolaou (2013, 2014), the IST risk premium estimates obtained through either the *Ishock* or *IMC* measures are negative, with the exception of the estimate obtained from the *IMC* measure after controlling for *TFP* (model (2c)), which is positive but statistically insignificant. The risk premia obtained from the *gIMC* measure are insignificant.

To assess the robustness of the above estimates, we repeat the above estimation on the earlier subsample ranging from 1930 to 1963. Panel B in Table 7 reports the risk premia estimates obtained over this sample period. Risk premia estimates for *Ishock*, and *IMC* are positive, although not statistically significant. The estimates from *gIMC* are mostly negative but insignificant. These estimates are in sharp contrast to the negative values obtained in Panel A. In Panel C we combine both subsamples and estimate risk premia over the entire 1930–2012 sample. Over this sample period, the risk premia estimates for *Ishock* and *IMC* are positive and statistically significant for at least three models (2a, 2c, and 2d). In contrast, estimates from *gIMC* are mostly negative and statistically insignificant.

Our analysis based on book-to-market portfolios as test assets suggests that the estimates of the IST risk premium is sensitive to both the sample period and the test assets. Estimates from only book-to-market portfolios are typically negative in the more recent

1964–2012 sample but are positive in both the earlier 1930–1963 sample and in the entire 1930–2012 sample. In contrast, as illustrated in the results of Table 2, our estimates of the IST risk premium using a set of 40 test assets are much less sensitive to the sample period used. This finding illustrates the importance of using a broad cross-section consisting of test assets sorted along different firm characteristics when estimating the IST risk premium.[9]

4.1.2 *IMC*-beta portfolios only

Kogan and Papanikolaou (2014) also use *IMC*-beta sorted portfolios as test assets and report significant negative IST risk premia estimates. For comparison, we also employ the ten *IMC*-beta sorted portfolios to estimate the IST risk premium. Table 8 reports the results. The risk premium estimates obtained over the 1930–2012 sample are indistinguishable from zero. Moreover, for the post-1963 sample, the IST risk premium estimates are sensitive to the econometric model specification. For example, when we allow for a constant term (the "Intercept") in the second-stage cross-sectional regressions, Panel B shows that the IST risk premium estimates are indistinguishable from zero. However, Panel C shows that the IST risk premium estimates become negative (except for the case where only *IMC* is used, model(2a)) and mostly significant if we restrict the intercept to be zero. Note that the intercept estimates in Panel B are all significantly different from zero (the only model for which we cannot reject a zero intercept is model (2b) with a t-stat of 1.61). This indicates that restricting the intercept to be zero as in Panel C can bias the estimates of the IST risk premium. As we will discuss below, the Fama-MacBeth approach without an intercept is equivalent to the GMM methodology used by Kogan and Papanikolaou (2014). This explains why they find a negative and significant IST risk premium, while our estimates, which are obtained by allowing for a nonzero intercept, are indistinguishable from zero.

[9] We also estimate the IST risk premium using only 30 industry portfolios. The estimates are mostly positive (but insignificant) for the 1930–2012 sample period and negative (but insignificant) for the 1964–2012 period. Results are available upon request.

4.2 Different econometric methodologies: Fama-MacBeth versus GMM

Our analysis relies on two-stage Fama-MacBeth regressions in estimating the IST risk premium. In principle, the Fama-MacBeth approach allows for flexibility in the choice of the sample used for estimating betas (e.g., full-window vs. rolling-window estimates).[10] However, the majority of studies that estimate the IST risk premium do so by using GMM to recover the price of risk parameter in a stochastic discount factor (SDF), where one of the risk sources is the IST shock. Theoretically, under the assumption that the regressors are not time-varying the estimates from the two methodologies should be identical and the difference should only concern the computation of the standard errors of the estimates (see Section 12.3 in Cochrane (2005)). In this subsection we verify that our analysis is indeed unaffected by whether we rely on two-stage Fama-MacBeth regressions or on GMM when estimating the IST risk premium.

The GMM approach usually starts by positing a model for the SDF, e.g.,

$$m = a - b_x \Delta x - b_z \Delta z, \tag{4}$$

where a is a constant, b_x and b_z are the prices of risk for the two shocks, x and z, respectively. In our setting, z is the capital-embodied IST shock and x is the disembodied shock (e.g., *MKT*, *TFP*, or gC), and Δx and Δz denote innovations to these shocks.

The model pricing errors are used as moment restrictions. That is, to estimate the parameters a, b_x, and b_z in (4) we require that the SDF m prices the cross-section of asset returns. In most applications, the SDF is normalized to one, i.e., $\mathbb{E}[m] = 1$, which allows to state the moment restrictions in terms of excess returns as follows:

$$\mathbb{E}[R_i^e] = -\text{cov}(m, R_i^e), \qquad \text{for all assets } i \tag{5}$$

[10]Note, however that the rolling-window approach is not suitable for our case because of the low frequency of our data.

where R_i^e denotes excess return over the risk-free rate of the i-th asset. In the estimation, moment restrictions are weighted by a weighting matrix (typically the identity matrix is used in the first-stage GMM estimate) and standard errors of the estimates are computed using the Newey and West (1987) procedure.

To compare the risk premia estimate from GMM to those from Fama-MacBeth, note that the price of risk parameters b_x and b_z in equation (4) are different from the IST risk premia λ_x and λ_z we computed in Section 3. The relation between these two quantities is given by (see, e.g., Section 13.4 of Cochrane (2005)):

$$\lambda = \mathbb{E}(ff')b, \tag{6}$$

where $\lambda = (\lambda_x, \lambda_z)'$, $f = (\Delta x, \Delta z)'$ and $b = (b_x, b_z)'$.

Table 9 reports the GMM estimates of the IST risk premium obtained from the GMM estimates of the prices of risk in (4) via the transformation (6). Panel A assumes that (5) holds, while Panel B allows a constant error in the pricing equation, i.e., $\mathbb{E}[R_i^e] + \mathrm{cov}(m, R_i^e) = \alpha$, for any asset i, where α is a constant to be estimated. Three points are worth of note. First, allowing for a constant pricing error in the moment conditions (5) is equivalent to allowing for an intercept in the second-stage Fama-MacBeth cross-sectional regressions. Indeed, the point estimates in Panel B of Table 9 are exactly identical to the point estimates in Panel A of Table 2.[11] The only difference between the two methodologies is in the t-statistics. However, the overall inference from the two approaches is the same. Second, when we impose the null hypothesis that the pricing error in (5) is zero (Panel A of Table 9), the point estimates of IST risk premium can be quite different from those obtained by allowing for a constant pricing error (Panel B). For example, in model (3a), the point estimate of IST risk premium is -11.5% assuming no pricing error (Panel A), but it becomes 2.09% if we allow pricing error (Panel B). Therefore, restricting the model to have a zero pricing error in GMM may bias the slope estimates. Finally, as we report in Table 1, the correlation between the IST measures (z)

[11] We have also verified that removing the intercept from the second-stage Fama-MacBeth cross-sectional regressions delivers the same point estimates as in Panel A of Table 9.

and the disembodied shocks ($x = MKT$, TFP, gC) may switch signs across different sample periods. Time-varying correlations across the factors may imply that the λ's and b's may have opposite signs when the factors are negatively correlated.

5 Discussion

In the above analysis we have provided empirical evidence that the pricing effect of IST shocks is sensitive to sample periods, test assets, and econometric methodologies. It is important to stress that these findings are obtained by using empirical proxies of IST shocks. From our discussion in Subsection 2.2, the data exhibit evident symptoms of potential measurement problems, emphasized by the low correlations among the three IST proxies documented in Table 1. In this section, we investigate further avenues to detect potential measurement problems in existing IST proxies and provide some suggestions on how future research efforts can address such issues.

Theoretically, IST shocks represent technological innovations that are embodied in new capital goods. The rationale behind *Ishock* as an IST proxy is that a better technology increases the supply of quality-adjusted capital goods, which leads to the decline in the relative price of capital goods. However, because the relative price of capital goods can also be affected by demand, *Ishock* is necessarily a noisy measure of IST shocks. The return-based measure, *IMC*, is motivated by the theoretical model of Papanikolaou (2011), in which neutral productivity shocks equally affect investment- and consumption-good firms while IST shocks only affect investment-good firms. Under this assumption, the return spread between investment- and consumption-goods producers is a factor-mimicking portfolio of IST shocks. The other measure, *gIMC*, captures the same idea but with macro quantities: a positive IST shock leads to an improvement in investment opportunities, so the aggregate investment increases more relative to the output of the consumption sector. The quality of both *IMC* and *gIMC* as proxies for IST shocks crucially depends on whether the structural assumptions of the model on which they rest are satisfied in the data.

One way to investigate potential measurement problems in IST proxies is to study the responses of macro quantities to these measures and compare them with theoretical predictions. Theory suggests that if the proxies are good measures of IST shocks, these theoretical arguments imply that aggregate consumption and investment should respond in a significant way to IST shocks. For example, the model of Papanikolaou (2011) predicts that investment increases and consumption decreases upon a positive IST shock. In contrast, the model of Garlappi and Song (2016a) predicts that both investment and consumption react positively to positive IST shocks.

To verify whether the IST proxies that we use in our study are related to macroeconomic aggregates, in Table 10 we compute the impulse responses of aggregate consumption and investment to IST shocks. We compute these impulse responses by regressing these quantities on each one of the three IST proxies. As Panel A of the table shows, in the annual 1930–2012 sample both consumption and investment react positively to *Ishock* and *gIMC*, consistent with the predictions of Garlappi and Song (2016a). On the contrary, consumption reacts negatively and investment reacts positively to *IMC*, although none of the coefficients are significant. According to the theory of Papanikolaou (2011) on which *IMC* is built, consumption should respond negatively to IST shocks and investment should respond positively. The fact that both consumption and investment do not seem to be related to *IMC* is somewhat problematic for this measure and suggests that the identifying structural assumptions that justify *IMC* as a proxy for IST shocks do not find strong support in the data.

Panel B of Table 10 further shows that in the more recent 1964–2012 annual sample, the impulse response of consumption to *Ishock* has lower significance than in the overall sample, suggesting that *Ishock* might be a more noisy measure in recent years. For quarterly data, Panel C of Table 10 shows that both investment and consumption react positively to *IMC*, although the coefficients are not statistically significant.

Overall, the impulse response analysis in Table 10 suggests that the two macro-based measures of IST shocks, *Ishock* and *gIMC*, can generate stronger comovement in the aggregate consumption and investment than the return-based measure *IMC*. This indicates that the measurement errors for the return-based measure are potentially larger than the macro-based measures.

The above evidence suggests that addressing measurement issues should be the primary focus of future research efforts. A potentially promising novel direction is to by-pass altogether the construction of IST proxies and attempt to measure directly firms' exposure to IST shocks from their observable investment activity. Building on this idea, Garlappi and Song (2016b) propose a new "model-free" measure of firms' exposure to investment shocks that, unlike existing proxy-based measures, can be computed directly from observable investment data and therefore less subject to the measurement issues discussed above. Based on post-1963 data, they find that value firms have higher investment-based IST exposures than growth firms, in contrast to the opposite pattern observed when using IST proxies. This demonstrates the importance of incorporating investment data in future research when investigating the pricing effect of IST shocks on cross-sectional equity returns.

6 Conclusion

In this paper we assess whether capital-embodied, investment-specific, technology shocks can explain the cross section of equity returns. We obtain three main results: (1) we find some weak evidence that IST shocks can explain value premium but much less so for the case of momentum profits and return spreads generated by asset growth, net share issues, earning-to-price ratio, and accrual; (2) using commonly used measures of IST shocks, a long data sample from 1930 to 2012, and a broad cross-section of 40 test assets, we estimate a *positive* risk premium for IST shocks; (3) we show that empirical inferences based on commonly used proxies of IST shocks are sensitive to the time period considered, the set of test assets employed, and the econometric model specification.

Our findings call for further efforts in understanding how investment shocks and heterogeneity in firms' investment decisions can generate cross-sectional return patterns of the magnitude observed in the data. In light of the measurement problems affecting IST proxies, exploring alternative measures of IST shocks to those existing in the literature appears to be of first-order importance to gain a better understanding of the effect of investment shocks on asset returns. Future research on the IST pricing effect should also explore the information contained in firms' investment data.

A Data details

A.1 Macroeconomic variables

Price deflator of consumption goods (P_C): the price deflator for nondurable consumption goods (row 5 of NIPA table 1.1.9). The annual series is available since 1929. The quarterly series is available since 1947I.

Price deflator of investment goods (P_I): the price deflator for equipment and software in the gross private domestic investment (row 11 of NIPA table 1.1.9). The availability of this series is the same as P_C. To take into account the quality adjustment, we employ instead the quality-adjusted series of Israelsen (2010).

Israelsen (2010) follows Gordon (1990) and Cummins and Violante (2002) and extends the annual quality-adjusted price series to the period of 1947–2006. We are grateful to Ryan Israelsen for kindly providing us with the long annual series for the period 1930–2012, which he constructed using the same methodology of Israelsen (2010).

Because the quarterly series of quality-adjusted investment goods price is not directly available, we approximate the growth rate of the quality-adjusted price from the unadjusted price. Specifically, we adjust equally the growth rates of investment good price for the four quarters in a year by the same amount as the annual quality adjustment. The annual growth rate adjustment is the difference in the growth rate between quality-adjusted price and NIPA's unadjusted price. This approach captures the year-to-year variation in quality adjustment while keeping the within year quarterly adjustment constant.

GDP and consumption expenditure growth: we measure economy-wide macroeconomic conditions using the annual growth rates of real GDP (row 1 of NIPA Table 1.1.1) and consumption (NIPA Table 1.1.1 contains personal consumption expenditures (PCE in row 2)). The annual data are available since 1930, and the quarterly data start from 1947II.

Growth rate spread in investment and consumption (gIMC): we measure the aggregate investment as the nonresidential investment (row 9 of NIPA Table 1.1.5) and the consumption as nondurable goods plus services (row 5 plus row 6 of NIPA Table 1.1.5). The $gIMC$ measure is the difference in the log growth rates of investment (gI) and consumption (gC). The annual data are available since 1930, and the quarterly data start from 1947II.

Total factor productivity (*TFP*): the annual total factor productivity data for 1930–1947 are from Kendrick (1961) (Table A-XXII for private domestic economy) and data for 1948–2012 are from the Bureau of Labor Statistics (multifactor productivity measure for private business sector). The quarterly data for 1947II–2012IV are from the Federal Reserve's business sector total factor productivity (available at http://www.frbsf.org/economic-research/total-factor-productivity-tfp/). In our regression analysis, we use the percentage change in TFP as a measure of the neutral technology risk.

A.2 Sector classification

Investment (I) and consumption (C) sectors: we rely on the procedure outlined in Gomes, Kogan, and Yogo (2009) and classify each Standard Industry Classification (SIC) code into either investment or consumption sector based on the 1987 benchmark input-output accounts. Gomes, Kogan, and Yogo (2009) provide a one-to-one match between SIC code and different categories of final demand, such as consumption (further classified as durable, nondurable, and services), investment, net exporter (NX), and government expenditure (G). Each industry specified by a SIC code is classified into the category of final demand to which it has the highest contribution. Their classification is available from Motohiro Yogo's website. We do not need the detailed classification within the consumption sector, and we allocate NX and G to either the investment or consumption sector depending on whether they contribute more to the investment or consumption sector.

A.3 Financial data

Return factors: the standard Fama-French 3 factors (*MKT*, *SMB*, *HML*) and the momentum (*UMD*) factors are all available at the monthly frequencies since January 1930 from Ken French's website. We then construct these factors at the quarterly and annual frequencies from the corresponding raw returns. For example, the annual market factor (*MKT*) is the raw annual market return minus the annual risk-free rate.

Test portfolios: we employ 10 size, 10 book-to-market, 10 momentum, and 10 industry portfolios in our main cross-sectional estimation of the IST risk premium. These portfolios are available starting from January 1930. In our robustness analysis, we also use an alternative set of 40 test portfolios, which include 10 profitability, 10 asset growth, 10 idiosyncratic volatility, and 10 net share issues. These portfolios are available only from July 1963. We construct the corresponding quarterly and annual series from the monthly portfolio returns, which are all downloaded from Ken French's website.

Standard cross-sections: We consider the effect of IST shocks on the following six return spreads that are generated by: (i) B/M, (ii) momentum, (iii) asset growth, (iv) net share issues, (v) E/P ratio, and (vi) accrual. All these cross-sections are downloaded from Ken French's website. Note that B/M and momentum portfolios are available since 1930, and all the other 4 cross-sections are available only from 1964.

***IMC* return:** to construct the *IMC* return, a firm's sector classification at June t is based on its SIC code from Compustat for the fiscal year ending in year $t-1$, if not missing, and on its SIC code from the Center for Research in Security Prices (CRSP) for June of year t, otherwise. The portfolio classification is then assigned to the firm for the next 12 months, from July of year t to June of year $t+1$. We calculate the value-weighted returns for each portfolio (I and C) using the lagged market value as weight, and then compound the monthly portfolio returns to quarterly and annual frequency based on calendar time. The *IMC* return is the investment sector return minus the consumption sector return. The *IMC* return is available starting from January 1930, at monthly, quarterly, and annual frequencies.

***IMC*-beta sorted portfolios:** to construct the *IMC*-beta sorted portfolios, at the end of each June, we sort firms into 10 value-weighted portfolios based on the past value of *IMC*-beta, which is estimated from a time series regression of weekly firm returns on weekly *IMC* returns for the previous 12 months. The portfolio ranking is then assigned to the firm in the next 12 months, from July of the sorting year to the June of the next year. Following Kogan and Papanikolaou (2014), we restrict the sample to firms producing consumption goods, and exclude financial firms.

Table 1: Time series properties of IST shocks

This table reports the time series properties of the three measures of investment-specific shocks over different sample periods. *Ishock* is based on the relative price of capital goods to consumption goods, as defined in equation (1). *IMC* is the return spread between firms in investment and consumption goods sectors defined in equation (2). *gIMC* is the growth rate difference in investment and consumption defined in equation (3). The reported summary statistics are in percentages (per year for the annual data and per quarter for quarterly data). Panels A, B, and C report results for *Ishock*, *IMC*, and *gIMC*, respectively. *PCE* is the growth rate of personal consumption expenditures, *GDP* is the growth rate of real gross domestic product, and *TFP* is the growth rate of total factor productivity. The return factors include Fama-French 3-factors (*MKT*, *SMB*, *HML*) and the momentum factor (*UMD*). The * and ** denote significance at the 10% and 5% levels, respectively.

			Correlations									
			Macro Factors			Return Factors				IST Measures		
	Mean	Std	*PCE*	*GDP*	*TFP*	*MKT*	*SMB*	*HML*	*UMD*	*Ishock*	*IMC*	*gIMC*
Panel A: *Ishock*												
Annual												
1930–2012:	3.45**	3.68	0.20*	0.46**	0.22**	0.07	0.15	0.11	0.26**	1.00	0.03	0.15
1930–1963:	2.00**	4.02	0.30*	0.68**	0.51**	0.50**	0.54**	0.43**	0.23	1.00	0.10	0.13
1964–2012:	4.46**	3.07	-0.01	0.21	-0.04	-0.35**	-0.19	-0.21	0.40**	1.00	0.02	0.34**
1948–2012:	3.87**	3.03	-0.02	0.16	-0.04	-0.29**	-0.05	-0.10	0.30**	1.00	0.03	0.23*
Quarterly												
1948–2012:	0.97**	1.25	0.00	0.14**	0.03	0.01	0.03	0.11*	0.03	1.00	0.12**	0.06
1964–2012:	1.11**	1.29	-0.01	0.20**	0.05	0.00	-0.04	0.08	0.07	1.00	0.14*	0.16**
Panel B: *IMC*												
Annual												
1930–2012:	0.61	14.17	0.04	-0.04	0.08	0.43**	0.23**	0.20*	0.13	0.03	1.00	0.12
1930–1963:	2.07	14.72	0.04	-0.09	0.02	0.60**	0.42**	0.69**	0.28	0.10	1.00	0.19
1964–2012:	-0.41	13.83	0.06	0.04	0.14	0.25*	0.10	-0.32**	0.04	0.02	1.00	-0.00
1948–2012:	0.03	12.93	0.09	0.05	0.16	0.35**	0.12	-0.11	0.06	0.03	1.00	0.00
Quarterly												
1948–2012:	0.02	5.20	0.20**	0.22**	0.25**	0.41**	0.27**	-0.13**	-0.00	0.12**	1.00	0.07
1964–2012:	-0.04	5.69	0.25**	0.23**	0.26**	0.39**	0.29**	-0.22**	-0.03	0.14*	1.00	0.07
Panel C: *gIMC*												
Annual												
1930–2012:	0.18	12.22	0.72**	0.16	0.28**	0.01	-0.12	0.01	0.26**	0.15	0.12	1.00
1930–1963:	0.41	18.12	0.79**	0.09	0.31*	0.06	-0.12	-0.01	0.30*	0.13	0.19	1.00
1964–2012:	0.02	5.41	0.47**	0.65**	0.21	-0.12	-0.18	0.10	0.42**	0.34**	-0.00	1.00
1948–2012:	0.14	5.48	0.50**	0.66**	0.28**	-0.21*	-0.22*	-0.02	0.45**	0.23*	0.00	1.00
Quarterly												
1948–2012:	0.03	2.32	0.29**	0.52**	0.29**	-0.10*	-0.12*	0.03	0.11*	0.06	0.07	1.00
1964–2012:	-0.01	1.95	0.30**	0.50**	0.22**	-0.10	-0.12*	0.04	0.08	0.16**	0.07	1.00

Table 2: Risk premium of IST shocks

This table reports the estimated IST risk premium (in percentage) from Fama-MacBeth cross-sectional regressions. The sample is based on annual data from 1930 to 2012 (Panel A) and 1964-2012 (Panel B), and the test assets are: size deciles, book-to-market deciles, momentum deciles, and 10 industry portfolios. The three IST measures are: *Ishock*, *IMC*, and *gIMC*. We consider both a one-factor model and two-factor models, with the second factor being either the market excess return (MKT), the growth rate of TFP (TFP), or the log growth rate of aggregate consumption (gC). The t-statistics in parentheses for the risk premium are adjusted for Shanken correction following Shanken (1992), and for autocorrelation and heteroskedasticity following Newey and West (1987).

	(1) *Ishock*				(2) *IMC*				(3) *gIMC*			
	(1a)	(1b)	(1c)	(1d)	(2a)	(2b)	(2c)	(2d)	(3a)	(3b)	(3c)	(3d)
Panel A: 1930-2012 sample												
Intercept	7.32	3.70	4.99	7.66	6.68	2.48	3.90	4.16	9.48	-1.32	3.79	7.30
	(2.59)	(0.82)	(1.68)	(2.44)	(3.49)	(0.55)	(1.10)	(1.52)	(4.35)	(-0.27)	(1.07)	(2.73)
λ_{Ishock}	3.77	3.31	3.08	3.81								
	(2.35)	(2.86)	(2.54)	(2.24)								
λ_{IMC}					3.83	1.98	0.55	3.01				
					(1.68)	(1.20)	(0.29)	(1.16)				
λ_{gIMC}									2.09	9.59	8.28	1.62
									(0.66)	(3.05)	(2.38)	(0.39)
λ_{MKT}		4.88				6.37				10.4		
		(1.09)				(1.33)				(2.00)		
λ_{TFP}			2.04				3.43				3.40	
			(1.16)				(1.98)				(1.80)	
λ_{gC}				0.97				4.71				4.40
				(0.45)				(2.56)				(2.16)
Adj.R^2	0.10	0.29	0.23	0.17	0.17	0.26	0.21	0.22	0.13	0.28	0.27	0.17
Panel B: 1964-2012 sample												
Intercept	8.37	7.44	8.03	7.91	7.39	8.36	7.06	9.36	8.52	3.72	7.88	9.60
	(3.99)	(1.88)	(3.34)	(2.61)	(4.16)	(2.10)	(3.38)	(4.06)	(4.30)	(0.74)	(3.25)	(4.11)
λ_{Ishock}	0.43	0.50	0.47	-1.06								
	(0.69)	(0.83)	(0.79)	(-0.96)								
λ_{IMC}					-0.14	0.17	-0.09	0.89				
					(-0.06)	(0.08)	(-0.04)	(0.34)				
λ_{gIMC}									1.91	2.94	2.44	0.94
									(1.60)	(2.03)	(2.22)	(0.57)
λ_{MKT}		0.11				-1.01				4.27		
		(0.03)				(-0.24)				(0.79)		
λ_{TFP}			0.30				0.25				0.62	
			(0.45)				(0.40)				(0.89)	
λ_{gC}				2.41				1.99				1.88
				(2.18)				(2.28)				(1.91)
Adj.R^2	0.14	0.28	0.25	0.24	0.15	0.27	0.26	0.25	0.15	0.27	0.25	0.26

Table 3: Portfolio returns and factor loadings on IST shocks: 1930–2012

This table reports the portfolio returns and their factor loadings on IST shocks. Panel A and B report quantities for portfolios sorted on book-to-market and past performance, respectively. The sample is based on annual data from 1930 to 2012. The average returns in excess of risk-free rate are in percentage per year. The factor loadings (betas) are estimated using the full-sample time series. We report the univariate loadings of portfolio returns on three IST measures. The column HML (WML) reports values for the high-minus-low (winner-minus-loser) portfolio. The t-statistics (in parentheses) are adjusted for autocorrelation and heteroskedasticity following Newey and West (1987).

Panel A: B/M portfolios

Variable	Low	2	3	4	5	6	7	8	9	High	HML
r	6.61	7.96	7.31	8.24	8.94	9.40	9.50	11.91	12.08	13.67	7.05
t-stat	(2.76)	(3.66)	(3.54)	(3.31)	(3.56)	(3.71)	(3.49)	(4.08)	(4.01)	(3.62)	(2.56)
β_{Ishock}	0.22	0.13	0.20	0.74	0.97	0.97	0.68	0.36	0.89	1.19	0.98
t-stat	(0.27)	(0.17)	(0.26)	(0.75)	(1.07)	(1.03)	(0.68)	(0.35)	(0.87)	(1.07)	(1.48)
β_{IMC}	0.66	0.46	0.46	0.71	0.65	0.69	0.64	0.83	0.73	1.09	0.43
t-stat	(4.47)	(3.86)	(3.94)	(3.10)	(3.06)	(3.56)	(2.47)	(2.90)	(2.43)	(2.81)	(1.06)
β_{gIMC}	-0.00	-0.09	-0.09	0.06	0.01	0.28	0.08	-0.01	-0.02	-0.23	-0.23
t-stat	(-0.00)	(-0.41)	(-0.44)	(0.24)	(0.02)	(1.04)	(0.27)	(-0.04)	(-0.05)	(-0.61)	(-1.06)

Panel B: Momentum portfolios

Variable	Loser	2	3	4	5	6	7	8	9	Winner	WML
r	0.80	5.25	5.60	7.20	6.68	7.79	8.60	10.41	11.47	15.70	14.90
t-stat	(0.22)	(1.77)	(2.16)	(2.82)	(3.01)	(3.32)	(3.82)	(4.40)	(4.53)	(5.36)	(5.47)
β_{Ishock}	-0.47	0.23	-0.02	0.46	0.04	0.10	0.31	0.90	0.96	1.22	1.69
t-stat	(-0.37)	(0.22)	(-0.02)	(0.51)	(0.05)	(0.11)	(0.35)	(1.16)	(1.19)	(1.26)	(1.80)
β_{IMC}	1.13	0.78	0.56	0.69	0.43	0.60	0.49	0.69	0.68	0.83	-0.29
t-stat	(4.21)	(4.08)	(2.94)	(2.76)	(2.96)	(3.29)	(3.08)	(3.85)	(3.62)	(4.68)	(-1.05)
β_{gIMC}	-0.52	-0.16	-0.11	-0.01	0.00	0.02	0.08	0.15	0.14	0.18	0.71
t-stat	(-1.45)	(-0.52)	(-0.38)	(-0.06)	(0.01)	(0.07)	(0.33)	(0.68)	(0.57)	(0.66)	(2.56)

Table 4: Expected value premium and momentum profits: 1930–2012

This table reports the estimated value premium and momentum profits (in percentages) based on the high-minus-low (HML) and winner-minus-loser (WML) portfolios. The sample is based on annual data from 1930 to 2012. The expected risk premium due to the exposure to a risk factor is calculated as the risk exposure (β) multiplied with the risk premium of the corresponding risk factor (λ). The risk exposures are estimated using the full sample as in Table 3, and the risk premia of risk factors are estimated using the same models as in Panel A of Table 2. Column t(diff1) reports the t-statistics testing the null hypothesis that the differences between the observed value premium or momentum profits and the expected values based on the IST exposure alone ($\beta_{IST}\lambda_{IST}$) are on average zero. Column t(diff2) reports the t-statistics testing the null hypothesis that the differences between the observed value premium or momentum profits and the expected values (using all the risk factors in the model) are on average zero. Note that for univariate models, the two tests are equivalent, and we therefore report only the first test. The t-statistics are adjusted for Shanken correction following Shanken (1992), and for autocorrelation and heteroskedasticity following Newey and West (1987).

Panel A: Value Premium

Factors	$\widehat{HML}$	$\frac{\widehat{HML}}{HML}$	t(diff2)	$\beta_{IST}\lambda_{IST}$	$\frac{\beta_{IST}\lambda_{IST}}{HML}$	t(diff1)
Ishock:	—	—	—	3.69	52%	0.95
Ishock & *MKT*:	4.79	68%	0.75	2.67	38%	1.32
Ishock & *TFP*:	4.28	61%	0.79	2.63	37%	1.39
Ishock & *gC*:	4.10	58%	0.83	4.34	62%	0.73
IMC:	—	—	—	1.64	23%	2.35
IMC & *MKT*:	2.85	40%	2.28	0.37	5%	2.50
IMC & *TFP*:	3.28	47%	0.90	0.23	3%	2.80
IMC & *gC*:	2.75	39%	1.26	1.33	19%	2.46
gIMC:	—	—	—	-0.47	-7%	3.24
gIMC & *MKT*:	2.36	33%	1.54	-2.27	-32%	2.74
gIMC & *TFP*:	2.54	36%	1.06	-2.60	-37%	3.29
gIMC & *gC*:	1.74	25%	1.70	-0.53	-8%	3.75

Panel B: Momentum Profit

Factors	$\widehat{WML}$	$\frac{\widehat{WML}}{WML}$	t(diff2)	$\beta_{IST}\lambda_{IST}$	$\frac{\beta_{IST}\lambda_{IST}}{WML}$	t(diff1)
Ishock:	—	—	—	6.37	43%	2.14
Ishock & *MKT*:	4.74	32%	3.33	5.90	40%	2.70
Ishock & *TFP*:	6.00	40%	2.27	4.95	33%	3.23
Ishock & *gC*:	5.07	34%	2.88	4.20	28%	2.68
IMC:	—	—	—	-1.12	-8%	6.21
IMC & *MKT*:	-1.40	-9%	6.24	-0.38	-3%	6.02
IMC & *TFP*:	3.86	26%	2.51	-0.17	-1%	6.38
IMC & *gC*:	4.87	33%	3.16	-0.71	-5%	6.15
gIMC:	—	—	—	1.47	10%	6.51
gIMC & *MKT*:	4.51	30%	3.70	6.82	46%	3.09
gIMC & *TFP*:	6.05	41%	2.26	5.82	39%	4.77
gIMC & *gC*:	3.60	24%	3.99	0.95	6%	7.19

Table 5: Expected cross-sectional return spreads: 1964–2012

This table reports the estimated expected cross-sectional return spreads (in percentages) for six cross-sections. The sample is based on annual data from 1964 to 2012. The expected risk premium due to the exposure to a risk factor is calculated as the risk exposure (β) multiplied with the risk premium of the corresponding risk factor (λ). The risk exposures are estimated using the 1964–2012 sample, and the risk premia of risk factors are estimated using the same models as in Panel B of Table 2. Column t(diff1) reports the t-statistics testing the null hypothesis that the differences between the observed return spreads and the expected return spreads based on the IST exposure alone ($\beta_{IST}\lambda_{IST}$) are on average zero. The t-statistics are adjusted for Shanken correction following Shanken (1992), and for autocorrelation and heteroskedasticity following Newey and West (1987).

Factors	$\beta_{IST}\lambda_{IST}$	$\frac{\beta_{IST}\lambda_{IST}}{HML}$	t(diff1)
Panel A: B/M (high minus low B/M)			
Ishock:	-0.66	-10%	2.89
IMC:	0.03	0.4%	2.30
gIMC:	-0.11	-1.6%	2.29
Panel B: Momentum (winners minus losers)			
Ishock:	1.13	7%	4.65
IMC:	0.05	0.3%	4.13
gIMC:	4.46	28%	4.22
Panel C: Asset Growth (low minus high growth)			
Ishock:	0.05	0.8%	2.61
IMC:	0.05	0.8%	2.51
gIMC:	-0.54	-9%	2.73
Panel D: Net Share Issues (low minus high issues)			
Ishock:	0.35	6%	2.18
IMC:	0.07	1.2%	2.63
gIMC:	0.01	0.2%	2.17
Panel E: Accrual (low minus high accrual)			
Ishock:	0.37	7%	3.62
IMC:	0.00	0.0%	4.14
gIMC:	0.64	11%	3.48
Panel F: E/P Ratio (high minus low E/P)			
Ishock:	-0.42	-7%	2.60
IMC:	0.09	1.4%	1.91
gIMC:	0.51	9%	1.92

Table 6: Risk premium of IST shocks—Alternative test assets

This table reports the estimated IST risk premium from Fama-MacBeth cross-sectional regressions. The table is the same as Table 2, except that the sample starts from 1964 instead of 1930 and the test assets are: (i) 10 profitability portfolios, (ii) 10 asset growth portfolios, (iii) 10 volatility portfolios, and (iv) 10 net share issues portfolios, discussed in Subsection 3.4. We use both the annual (Panel A) and quarterly (Panel B) data, and the full sample is used in the first-stage beta estimation. The t-statistics in parentheses for the risk premium are adjusted for Shanken correction following Shanken (1992), and for autocorrelation and heteroskedasticity following Newey and West (1987).

	(1) *Ishock*				(2) *IMC*				(3) *gIMC*			
	(1a)	(1b)	(1c)	(1d)	(2a)	(2b)	(2c)	(2d)	(3a)	(3b)	(3c)	(3d)
Panel A: annual data												
Intercept	9.26	9.99	9.62	10.1	7.20	5.01	7.19	7.27	6.03	9.43	6.82	6.73
	(3.10)	(2.32)	(2.93)	(2.90)	(3.79)	(1.01)	(3.58)	(3.59)	(2.80)	(1.90)	(2.43)	(2.64)
λ_{Ishock}	1.32	1.05	1.25	1.27								
	(0.98)	(1.10)	(0.96)	(0.94)								
λ_{IMC}					-1.84	-2.81	-1.84	-1.80				
					(-0.77)	(-1.04)	(-0.78)	(-0.70)				
λ_{gIMC}									-0.83	-3.92	-1.59	-1.83
									(-0.26)	(-1.95)	(-0.57)	(-0.55)
λ_{MKT}		-3.37				1.59				-3.11		
		(-0.78)				(0.29)				(-0.66)		
λ_{TFP}			-0.41				-0.02				-0.71	
			(-0.53)				(-0.04)				(-1.17)	
λ_{gC}				0.75				0.11				0.75
				(0.73)				(0.12)				(0.68)
Adj.R^2	0.12	0.33	0.18	0.23	0.30	0.33	0.31	0.34	0.12	0.32	0.18	0.19
Panel B: quarterly data												
Intercept	1.50	2.20	1.85	1.50	1.86	1.76	1.76	1.87	1.47	2.17	1.56	1.50
	(1.92)	(1.93)	(2.50)	(2.09)	(3.49)	(1.77)	(3.14)	(3.21)	(2.45)	(2.02)	(2.13)	(2.52)
λ_{Ishock}	1.05	1.05	1.11	1.05								
	(3.58)	(3.58)	(3.17)	(3.72)								
λ_{IMC}					-0.45	-0.51	-0.52	-0.44				
					(-0.69)	(-0.77)	(-0.77)	(-0.63)				
λ_{gIMC}									-0.17	-1.27	-1.12	0.01
									(-0.16)	(-1.91)	(-1.67)	(0.01)
λ_{MKT}		-0.63				-0.18				-0.61		
		(-0.54)				(-0.16)				(-0.54)		
λ_{TFP}			-0.54				0.64				-1.32	
			(-0.42)				(0.91)				(-1.18)	
λ_{gC}				0.17				0.32				0.07
				(0.48)				(1.51)				(0.36)
Adj.R^2	0.00	0.31	0.27	0.15	0.30	0.31	0.31	0.34	0.15	0.33	0.29	0.20

Table 7: Risk premium of IST shocks: Estimates from B/M portfolios

This table reports the estimated IST risk premium (in percentages) from Fama-MacBeth cross-sectional regressions based on 10 B/M portfolios. The estimation methods used are the same as those described in Table 2 with the following exceptions: (i) the test assets are the 10 book-to-market portfolios; (ii) we consider sub-samples of the period 1930–2012. The t-statistics for the risk premium are adjusted for Shanken correction following Shanken (1992), and for autocorrelation and heteroskedasticity following Newey and West (1987).

	(1) *Ishock*				(2) *IMC*				(3) *gIMC*			
	(1a)	(1b)	(1c)	(1d)	(2a)	(2b)	(2c)	(2d)	(3a)	(3b)	(3c)	(3d)
Panel A: 1964-2012 sub-sample												
Intercept	0.79	3.55	1.33	1.70	8.79	-14.5	4.54	8.80	9.51	3.24	3.26	9.12
	(0.20)	(0.58)	(0.37)	(0.49)	(4.09)	(-1.42)	(1.58)	(3.97)	(3.25)	(0.42)	(0.88)	(3.62)
λ_{Ishock}	-3.09	-3.02	-2.12	-2.93								
	(-2.39)	(-2.56)	(-2.20)	(-2.65)								
λ_{IMC}					-5.92	-13.6	0.99	-5.81				
					(-1.39)	(-1.42)	(0.31)	(-2.42)				
λ_{gIMC}									4.10	7.11	-0.32	1.91
									(1.07)	(1.36)	(-0.10)	(0.73)
λ_{MKT}		3.26				21.3				5.25		
		(0.53)				(2.29)				(0.69)		
λ_{TFP}			1.00				1.80				1.88	
			(0.90)				(1.91)				(1.58)	
λ_{gc}				0.59				0.16				1.24
				(0.46)				(0.12)				(1.11)
Adj.R^2	0.21	0.32	0.36	0.38	0.14	0.25	0.36	0.12	0.04	0.05	0.32	0.16
Panel B: 1930-1963 sub-sample												
Intercept	4.61	1.13	5.18	5.68	5.32	-1.18	5.22	5.89	12.6	1.34	9.99	4.06
	(1.23)	(0.22)	(1.50)	(2.23)	(1.57)	(-0.14)	(1.62)	(2.48)	(2.68)	(0.34)	(2.72)	(0.75)
λ_{Ishock}	2.12	0.63	1.82	1.51								
	(1.55)	(0.64)	(1.39)	(1.34)								
λ_{IMC}					5.30	1.90	5.27	5.79				
					(1.71)	(0.89)	(1.75)	(1.50)				
λ_{gIMC}									-5.15	0.48	-3.71	-5.15
									(-0.78)	(0.09)	(-0.58)	(-0.52)
λ_{MKT}		9.41				11.9				9.22		
		(1.81)				(1.44)				(1.80)		
λ_{TFP}			0.17				0.11				1.30	
			(0.20)				(0.13)				(1.01)	
λ_{gc}				0.18				-1.40				7.31
				(0.07)				(-0.45)				(0.95)
Adj.R^2	0.30	0.37	0.31	0.39	0.33	0.37	0.34	0.39	0.09	0.44	0.17	0.30
Panel C: 1930-2012 full-sample												
Intercept	7.17	-5.24	5.15	7.53	2.35	-8.56	2.30	2.23	9.56	-5.26	4.40	4.18
	(2.78)	(-1.04)	(2.09)	(2.78)	(0.85)	(-1.13)	(0.86)	(0.83)	(4.03)	(-1.05)	(1.59)	(0.87)
λ_{Ishock}	3.76	0.05	2.66	3.97								
	(1.77)	(0.04)	(1.41)	(1.69)								
λ_{IMC}					10.4	-0.89	9.63	9.83				
					(2.47)	(-0.28)	(2.23)	(2.41)				
λ_{gIMC}									-5.08	0.67	-3.18	-7.40
									(-1.68)	(0.22)	(-0.86)	(-1.00)
λ_{MKT}		13.8				17.0				13.8		
		(2.79)				(2.27)				(2.56)		
λ_{TFP}			2.05				0.55				2.50	
			(2.05)				(0.68)				(1.92)	
λ_{gc}				0.91				0.27				7.98
				(0.39)				(0.15)				(1.10)
Adj.R^2	0.15	0.34	0.17	0.20	0.22	0.30	0.18	0.29	0.05	0.36	0.22	0.35

Table 8: Risk premium of IST shocks: Estimates from *IMC*-beta portfolios

This table reports the estimated IST risk premium (in percentages) from Fama-MacBeth cross-sectional regressions based on 10 *IMC*-beta sorted portfolios. The estimation methods used are the same as those described in Table 7. The t-statistics for the risk premium are adjusted for Shanken correction following Shanken (1992), and for autocorrelation and heteroskedasticity following Newey and West (1987).

	(1) *Ishock*				(2) *IMC*				(3) *gIMC*			
	(1a)	(1b)	(1c)	(1d)	(2a)	(2b)	(2c)	(2d)	(3a)	(3b)	(3c)	(3d)
Panel A: 1930–2012 full-sample												
Intercept	9.05	8.66	9.05	9.17	8.68	8.41	9.09	8.92	8.77	8.44	8.90	8.91
	(4.59)	(2.70)	(4.49)	(4.28)	(4.73)	(1.21)	(4.52)	(4.28)	(4.40)	(2.70)	(4.18)	(3.57)
λ_{Ishock}	-0.50	-0.65	-0.50	-0.62								
	(-0.27)	(-0.43)	(-0.30)	(-0.40)								
λ_{IMC}					0.16	0.08	0.44	0.09				
					(0.10)	(0.04)	(0.21)	(0.06)				
λ_{gIMC}									-0.14	-0.21	0.01	-0.09
									(-0.05)	(-0.06)	(0.00)	(-0.03)
λ_{MKT}		0.23				0.36				0.33		
		(0.07)				(0.05)				(0.10)		
λ_{TFP}			-0.08				-0.38				-0.07	
			(-0.09)				(-0.39)				(-0.08)	
λ_{g_c}				-0.65				-0.34				-0.30
				(-0.41)				(-0.20)				(-0.16)
Adj.R^2	-0.04	0.27	0.10	0.14	0.28	0.28	0.27	0.28	-0.02	0.29	0.13	0.28
Panel B: 1964–2012 sub-sample												
Intercept	10.2	11.0	10.6	10.8	7.23	7.77	7.22	7.33	6.75	7.42	6.87	7.15
	(2.37)	(2.62)	(2.83)	(3.02)	(4.08)	(1.61)	(3.64)	(2.84)	(3.13)	(2.03)	(2.95)	(2.94)
λ_{Ishock}	1.46	1.55	1.58	1.55								
	(0.82)	(0.96)	(1.00)	(0.97)								
λ_{IMC}					-0.36	-0.15	-0.36	-0.31				
					(-0.17)	(-0.05)	(-0.17)	(-0.15)				
λ_{gIMC}									-0.77	-0.79	-0.81	-0.90
									(-0.43)	(-0.44)	(-0.46)	(-0.50)
λ_{MKT}		-3.83				-0.66				-0.42		
		(-0.86)				(-0.12)				(-0.10)		
λ_{TFP}			-0.22				0.01				-0.15	
			(-0.24)				(0.01)				(-0.18)	
λ_{gC}				0.30				0.08				0.20
				(0.20)				(0.06)				(0.15)
Adj.R^2	-0.06	0.28	0.02	0.16	0.29	0.33	0.28	0.36	0.01	0.35	0.10	0.24
Panel C: 1964–2012 sub-sample, no intercept												
λ_{Ishock}	-3.42	-2.93	-3.34	-3.20								
	(-2.75)	(-1.66)	(-2.74)	(-2.15)								
λ_{IMC}					7.14	-3.05	2.16	-2.82				
					(2.26)	(-1.34)	(0.45)	(-1.00)				
λ_{gIMC}									-11.6	-3.88	-8.32	-7.88
									(-1.69)	(-1.65)	(-1.87)	(-1.93)
λ_{MKT}		7.03				7.14				6.76		
		(5.67)				(3.99)				(4.23)		
λ_{TFP}			0.26				5.53				0.85	
			(0.22)				(1.23)				(0.56)	
λ_{gC}				-0.35				-4.94				-3.00
				(-0.19)				(-1.85)				(-1.42)
Adj.R^2	0.61	0.73	0.64	0.69	0.23	0.75	0.48	0.72	0.54	0.66	0.58	0.64

Table 9: Risk premium of IST shocks: GMM

This table reports the estimated IST risk premium (in percentage) from GMM approach. The sample is based on annual data from 1930 to 2012, and the test assets are: size deciles, book-to-market deciles, momentum deciles, and 10 industry portfolios. The three IST measures are: *Ishock*, *IMC*, and *gIMC*. We consider both a one-factor model and two-factor models, with the second factor being either the market return (MKT), the growth rate of TFP (TFP), or the log growth rate of aggregate consumption (gC). We report the estimates from first-stage GMM using identity weighting matrix. We also report the mean absolute pricing error (MAPE) for each model. The t-statistics for the risk premium are adjusted for autocorrelation and heteroskedasticity following Newey and West (1987).

	(1) *Ishock*				(2) *IMC*				(3) *gIMC*			
	(1a)	(1b)	(1c)	(1d)	(2a)	(2b)	(2c)	(2d)	(3a)	(3b)	(3c)	(3d)
Panel A: assuming zero error in moments												
λ_{Ishock}	10.8	3.14	2.80	9.12								
	(0.88)	(3.16)	(1.48)	(0.97)								
λ_{IMC}					11.6	1.22	1.36	5.51				
					(1.88)	(0.77)	(0.31)	(1.13)				
λ_{gIMC}									-11.5	8.76	11.4	-1.78
									(-1.55)	(2.00)	(1.30)	(-0.08)
λ_{MKT}		8.18				8.68				9.16		
		(2.04)				(3.55)				(2.27)		
λ_{TFP}			4.73				5.40				5.32	
			(1.62)				(1.80)				(1.66)	
λ_{gC}				13.0				7.62				14.5
				(0.70)				(1.43)				(0.86)
MAPE (%)	4.24	1.45	1.89	3.64	2.70	1.64	1.91	1.99	9.26	1.73	1.92	3.22
Panel B: assuming a constant error in moments												
Constant	7.32	3.70	4.99	7.66	6.68	2.48	3.90	4.16	9.48	-1.32	3.79	7.30
	(1.64)	(1.10)	(1.53)	(1.67)	(3.06)	(0.65)	(1.13)	(1.03)	(4.18)	(-0.36)	(1.01)	(1.79)
λ_{Ishock}	3.77	3.31	3.08	3.81								
	(3.03)	(3.16)	(2.44)	(3.13)								
λ_{IMC}					3.83	1.98	0.55	3.01				
					(1.95)	(1.57)	(0.16)	(1.11)				
λ_{gIMC}									2.09	9.59	8.28	1.62
									(0.59)	(2.44)	(1.58)	(0.25)
λ_{MKT}		4.88				6.37				10.4		
		(1.29)				(1.62)				(2.10)		
λ_{TFP}			2.04				3.43				3.40	
			(1.39)				(2.54)				(2.21)	
λ_{gC}				0.97				4.71				4.40
				(0.55)				(2.64)				(2.41)
MAPE (%)	1.61	1.43	1.48	1.57	1.80	1.68	1.66	1.75	2.21	1.72	1.69	2.12

Table 10: Impulse responses of consumption and investment to IST shocks

This table reports the result from univariate regressions of aggregate consumption and investment on the three proxies of IST shocks: *Ishock*, *IMC*, and *gIMC*. The log growth rates of aggregate consumption and investment are the same as those used to construct *gIMC* measure (see Appendix A for details). The t-statistics (in parenthesis) are adjusted for autocorrelation and heteroskedasticity following Newey and West (1987).

	Consumption			Investment		
Panel A: annual sample of 1930–2012						
Intercept	0.037	0.060	0.060	0.022	0.062	0.060
	(2.17)	(7.35)	(8.12)	(0.45)	(2.84)	(8.12)
Ishock	0.655			1.164		
	(2.04)			(1.21)		
IMC		-0.047			0.058	
		(-1.12)			(0.52)	
gIMC			0.198			1.198
			(1.48)			(8.96)
Adj.R^2	0.214	0.005	0.215	0.067	-0.009	0.913
Panel B: annual sample of 1964–2012						
Intercept	0.069	0.069	0.069	0.043	0.070	0.069
	(8.00)	(12.40)	(13.18)	(2.22)	(5.68)	(13.18)
Ishock	0.016			0.606		
	(0.13)			(2.41)		
IMC		-0.022			-0.022	
		(-0.83)			(-0.26)	
gIMC			0.167			1.167
			(1.88)			(13.17)
Adj.R^2	-0.021	-0.006	0.115	0.057	-0.019	0.880
Panel C: quarterly sample of 1964Q1–2012Q4						
Intercept	0.016	0.017	0.017	0.013	0.017	0.017
	(10.69)	(21.42)	(22.30)	(4.15)	(7.46)	(22.30)
Ishock	0.154			0.395		
	(1.73)			(2.47)		
IMC		0.013			0.036	
		(1.43)			(1.12)	
gIMC			0.076			1.076
			(1.51)			(21.29)
Adj.R^2	0.069	0.005	0.036	0.048	0.003	0.894

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