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Trade policy uncertainty and stock returns ☆

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ABSTRACT

A recent literature has documented large real effects of trade policy uncertainty (TPU) on trade, employment, and investment, but there is little evidence that investors are compensated for bearing such risk. To quantify the risk premium associated with TPU, we exploit quasi-experimental variation in exposure to TPU arising from Congressional votes to revoke China's preferential tariff treatment between 1990 and 2001. A long-short portfolio designed to isolate exposure to TPU earns a risk-adjusted return of 3.6–6.2% per year. This effect is larger in sectors less protected from globalization, and more reliant on inputs from China. Industries more exposed to trade policy uncertainty also had a larger drop in stock prices when the uncertainty began, and more volatile returns around key policy dates. Our results are not explained by the effects of policy uncertainty on expected cash-flows, investors' forecast errors, and import competition from China.

1. Introduction

The recent trade war between the U.S. and China has brought trade policy uncertainty (henceforth TPU) to the forefront of the economic and policy debate. A growing empirical literature has analyzed the effects of TPU on employment (Pierce and Schott (2016)), international trade (Handley and Limão (2017)), and investment (Pierce and Schott (2018), Caldara et al. (2020)). While this literature has focused on the impact of TPU on economic outcomes, it has remained silent on its effect on asset prices. Given the relevance that stock prices have for firms' investment decisions (Chen et al. (2006)), household wealth (Guiso et al. (2002)), and employment (Chodorow-Reich et al. (2019)), studying how financial markets respond to policy uncertainty deserves further scrutiny.

This paper documents that TPU is a systematic risk factor that affects asset prices. We capture industries' exogenous and heterogeneous exposure to TPU arising from Congressional votes to revoke China's preferential tariff rates between 1990 and 2001, before China's accession to the World Trade Organization (WTO). Empirically, we find a large risk premium associated with exposure to TPU: investors required an additional 3.6–6.2% return per year on average as compensation for uncertainty about future trade policy. Moreover, the risk premium was larger in sectors more exposed to globalization and more reliant on inputs from China.

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We focus on the uncertainty arising from annual votes by Congress to revoke China's "Most Favored Nation" (MFN) status between 1990 and 2001. Starting in 1980, U.S. imports from China were subject to Normal Trade Relations (NTR), or equivalently MFN, tariff rates reserved for WTO members, even though China was not a member of the WTO. This required annual renewals by Congress, which were essentially automatic until the Tiananmen Square Massacre in 1989. Starting in 1990, NTR renewal in Congress became more politically contentious, with the House passing resolutions against Chinese NTR renewal in 1990, 1991 and 1992. China's tariff status, however, did not change because the Senate did not pass the House resolutions. Had NTR status been revoked, tariffs would have reverted to non-NTR rates, established under the Smoot-Hawley Tariff Act of 1930, which were on average 27% higher than existing tariffs. The uncertainty about tariff increases on Chinese goods ended on 12/11/2001, when China joined the WTO, eliminating the need for annual renewal votes.

We follow Pierce and Schott (2016), and quantify the heterogeneous exposure to policy uncertainty via the "NTR gap," defined as the difference between the non-NTR rates, to which tariffs would have risen if annual renewal had failed, and the NTR rates. We exploit the large cross-sectional variation in the NTR gaps across U.S. tradeable sectors and estimate a differences-in-differences specification in which we regress monthly firm level stock returns, between 1980 and 2001, on the industry-level NTR gap, interacted with a dummy for the period of trade policy uncertainty (hereafter TPU period). The identification relies on the fact that most of the cross-sectional variation in the NTR gaps comes from variation in non-NTR rates, set by the Smoot-Hawley Act in 1930, which are likely exogenous to the U.S. industries' stock returns 70 years after they were set.

Our baseline results suggest that U.S. firms more exposed to tariff uncertainty, i.e. firms operating in industries that had a higher gap between non-NTR and NTR rates, experienced significantly higher stock returns than less exposed firms between 1990 and 2001. We exploit the granularity of our dataset, which consists of 5,655 firms in 281 tradeable industries over 252 months, and control for unobserved firm- and time-specific characteristics, for time varying firm financial fundamentals, and other trade policy changes, such as the global Multi-Fiber Arrangement (MFA), the reduction in Chinese import tariffs associated with China's accession to WTO, as well as the exposure to NAFTA. The difference in average returns for high and low NTR gap firms is significant at the 1% level, and implies that going from an industry less exposed to TPU (at the 25th percentile of the distribution of NTR gaps in 1990), to an industry more exposed (at the 75th percentile of the distribution), would increase stock returns by 6.8% per year during the uncertainty period. We obtain similar results when we construct a firm-level exposure to TPU, by interacting the industry-level NTR gap with the share of the firm's sales from each major business segment, computed using Compustat Segments data. When we estimate a dynamic version of our baseline regression, we find that the effect of the NTR gap on stock returns was larger in the early 1990s. This is consistent with greater uncertainty about China's NTR status between 1990 and 1992, because this is when the House of Representatives passed resolutions revoking China's preferential tariff treatment.

We then argue that the higher average returns earned by more exposed sectors can be explained by a risk premium for exposure to TPU. Intuitively, during the TPU period, investors required additional compensation to hold stocks with exposure to this policy uncertainty, as argued in the theoretical frameworks of Pastor and Veronesi (2012) and Pástor and Veronesi (2013). Using the predictions of these models to guide our empirical analysis, we document several stylized facts in support of the risk premium hypothesis.

First, we show that a value-weighted portfolio, long high-gap firms and short low-gap firms, had average excess returns of 6.2% per year, over the 5 factors in Fama and French (2015), during the uncertainty period 1990–2001. Our long-short portfolio's excess return remains economically large and statistically significant after controlling for a firm's exposure to the text-based measures of economic and trade policy uncertainty constructed in Baker et al. (2016). When we form the portfolios using the firm-level exposure measure computed from Compustat Segments data, we find a similar risk premium, although the point estimate is smaller than the baseline. Our results suggest that, among the set of industries we consider, TPU was a systematic factor that could not be diversified away. Reassuringly, when we repeat the portfolio analysis in the control periods 1980–1989 and 2002–2007, and for a set of foreign countries in a placebo exercise, we find no significant difference in stock returns between high and low NTR gap industries.

Second, we investigate which firms, among the high-gap industries, are driving the estimated risk premium. First, we double sort portfolios on the NTR gap in 1990 and on NTR tariff rates. The risk premium was concentrated in industries whose exposure to globalization, as measured by low NTR tariff rates, was higher. We find similar results when we double sort portfolios on the NTR gap and on industry-level shipping costs from Barrot et al. (2018). This suggests that exposure to globalization amplifies the effects of trade policy uncertainty and thus the associated risk premium. An additional amplification channel is the share of inputs sourced from China: the risk premium was larger among firms with a higher share of inputs expenditures from China, suggesting that uncertainty about the cost of production was also priced by financial markets.[1]

Third, we document a large and significant drop in stock prices for industries with higher NTR gaps around the day in which the policy uncertainty reasonably began, i.e., the day when the first resolution to revoke China's NTR status was proposed in the House, on 07/23/1990. This is consistent with an increase in the discount rate due to higher uncertainty on future policies that depressed prices of more exposed industries, as in Pastor and Veronesi (2012).

[1] In this respect, we contribute to the literature that shows the importance of input–output linkages for economic outcomes (see e.g., Caliendo and Parro (2014), Wang et al. (2018), Adao et al. (2020), Esposito and Hassan (2021) and Blaum et al. (2020)), and for stock returns (see e.g., Cohen and Frazzini (2008) and Huang et al. (2018)).

Fourth, we document that more exposed firms had significantly larger realized volatility around the Congressional votes to revoke China's NTR status, consistent with the evidence shown, in other contexts, e.g., Boutchkova et al. (2012) and Baker et al. (2019).

To lend further support to our risk premium hypothesis, we discuss three other potential explanations for our results and show that they are not supported by the data. First, it could be that the differential returns of high and low NTR gap industries may have been driven by stock prices' responses to changes in expected cash-flows (as shown in e.g., Greenland et al. (2019)), instead of a risk premium. To disentangle the two effects, we first repeat the portfolio analysis, but exclude a 3-day window around the dates Permanent Normal Trade Relations (PNTR) status was granted, and around House votes to revoke China's NTR status. We find that removing these days does not significantly alter the estimated risk premium, although the point estimate is smaller than the baseline. In addition, we explicitly look at the returns of firms that were hurt the most after the trade policy uncertainty was resolved. To this end, we interact the difference in EBIT growth between 1990–2001 and 2002–2007 with our baseline DID term. If the investors knew that high gap firms would have been economically hurt by the resolution of uncertainty (as shown ex-post by Pierce and Schott (2016)), the interaction term should be negative and significant. Instead, we find no difference in average returns between firms that ex-post did poorly and ex-post did well after PNTR went into effect. We also find that including this interaction term does not qualitatively change our coefficient of interest.

Second, it could be that investors initially either under- or over-estimated the effect of TPU on firms' performance. If this were true, we would expect to find large stock-market responses around the dates U.S. firms released fundamental information in their quarterly earnings announcements. However, we find no evidence of systematic differences in earnings announcement returns for stocks with low and high NTR gaps. A third explanation for our results could be that more uncertainty about future tariffs offered implicit protection against Chinese competition for firms in high-gap industries, which, as a result, enjoyed higher cash flows and higher stock returns during the 1990–2001 period. We test for this channel by double sorting portfolios on NTR gap and import penetration from China, and find that, instead, the risk premium was earned by industries *more* exposed to China. This is consistent with our previous results that the risk premium estimates were driven by sectors more vulnerable to globalization.

Our paper complements the empirical literature that investigates the effects of trade policy uncertainty, and uncertainty in general, on economic outcomes, such as Pierce and Schott (2016), Handley and Limão (2017), Crowley et al. (2018), Chinn et al. (2018), Steinberg (2019), Esposito (2020), and Caldara et al. (2020). Differently from this literature, we focus on how TPU affects the riskiness perceived by investors, tracing down its effects on firms' stock returns. Stock returns are an important determinant of real economic variables (see e.g., Chodorow-Reich et al. (2019) and Jordà et al. (2019)), and thus their large and heterogenous response to TPU, documented in this paper, may have exacerbated the effects of TPU on economic outcomes.

There is an extensive literature that attempts to empirically assess how policy uncertainty is priced into stocks and options (see Pastor and Veronesi (2012), Pástor and Veronesi (2013), Brogaard and Detzel (2015), Kelly et al. (2016) and Bali et al. (2017)), but the intrinsic endogeneity of policy actions makes it difficult to identify the causal effects of policy uncertainty. The methodology used in this paper presents some advantages relative to this literature. First, the identification strategy relies on non-NTR tariff rates that were set 70 years before the onset of policy uncertainty, providing the quasi-experimental variation needed to estimate the risk premium. Second, while most indicators of policy uncertainty used by the literature do not vary across industries, see e.g., Jurado et al. (2015), Baker et al. (2016), our measure directly captures differences in exposure to TPU across sectors, and even at the firm level when we use Compustat Segments data.[2] Third, it is an *ex-ante* measure of uncertainty, and thus is not subject to a look-ahead bias.[3] Lastly, it is directly observable, and thus its construction is not subject to measurement error, and it does not rely on assumptions about the underlying volatility process.

We contribute to the literature on the effects of globalization on stock returns, see e.g., Fillat and Garetto (2015) and Barrot et al. (2018). This literature has shown that industries more exposed to foreign competition or foreign shocks command a large risk premium. Our double sorting exercise documents that the interaction between such "first-moment" effects are amplified by the "second-moment" effect arising from TPU.

There is a recent literature that uses stock-market event studies to evaluate trade policies (see e.g., Breinlich (2014), Moser and Rose (2014), Huang et al. (2018), Greenland et al. (2019)). The goal of this literature is to look at the short-run response of stock returns to policy news in order to tease out the market expectations on future cash-flows. Our goal is complementary, in that we look at the long-run behavior of stock returns, which is informative of how investors perceive firms' riskiness.

The paper proceeds as follows. Section 2 documents the effect of tariff uncertainty on average stock returns across U.S. tradeable industries during the 1990–2001 period. Section 3 argues that such effect was a risk premium for exposure to trade policy uncertainty. Section 4 discusses alternative explanations for the results. Section 5 concludes.

[2] Brogaard and Detzel (2015) use the Baker et al. (2016) index, which does not vary across industries, but allow firms to heterogeneously load on this factor in firm-level regressions. Caldara et al. (2020) use transcripts from firms' earnings conference calls to construct a firm-level measure of exposure to TPU.

[3] For instance, the widely-used method in Carr and Wu (2008) uses ex-post realized variance as a proxy for ex-ante expected variance, introducing a look-ahead bias. Kelly et al. (2016) and Alfaro et al. (2018) use forward-looking option-implied volatilities and realized volatilities.

2. Tariff uncertainty and U.S. stock returns

In this section, we use quasi-exogenous variation in exposure to tariff uncertainty across U.S. tradeable industries to identify the causal effect of trade policy uncertainty on stock returns. We first look at a difference-in-differences specification, which shows that firms in industries more exposed to TPU between 1990–2001 had higher average returns than low-gap industries, relative to the previous decade. We use this diff-in-diff result to motivate the portfolio analysis that we perform in Section 3 using only the uncertainty period 1990–2001.

2.1. Data and identification strategy

Starting in 1980, U.S. imports from China were subject to the relatively low Normal Trade Relations (NTR) tariff rates reserved for members of the World Trade Organization (WTO).[4] From 1980 to 1989, renewal of these NTR rates for China was essentially automatic. After the Tiananmen Square Massacre in 1989, however, the U.S. House of Representatives introduced and voted on legislation to revoke China's temporary NTR tariffs every year from 1990 to 2001. If Congress had failed to roll over the NTR rates, import tariffs on Chinese goods would have reset to the higher rates established in the Smoot-Hawley Tariff Act of 1930. The renewal process was politically contentious. In fact, as shown in Table A.1 in Appendix A, the House passed resolutions against Chinese NTR renewal in 1990, 1991 and 1992, despite being disapproved by the Senate later on.[5] In October 2000, the United States granted China Permanent Normal Trade Relations (PNTR) conditional on China joining the WTO. China joined the WTO at the end of 2001, and PNTR went into effect at the start of 2002. Granting China PNTR permanently removed this source of uncertainty by fixing U.S. tariffs on Chinese goods at NTR levels.

We argue that the annual Congressional votes generated uncertainty because: (i) investors were uncertain about whether China's NTR status would be revoked, and (ii) they were uncertain about the future performance of U.S. industries if NTR status was revoked. As in Pastor and Veronesi (2012), we refer to the former type of uncertainty as "political uncertainty", while to the latter as "policy uncertainty". While the likelihood of a policy change – the political uncertainty – was the same for all industries, since either all would revert to Smoot-Hawley rates, or all would keep lower NTR rates, the potential impact of the policy change – the policy uncertainty – could have been different across industries and firms.

To capture the exposure of each sector to such policy uncertainty, we follow Pierce and Schott (2016) and construct the "NTR gap", defined as the difference between the NTR and non-NTR rates to which tariffs would have risen if annual renewal had failed:

$$NTRGap_{it} = NonNTR_i - NTR_{it} \tag{1}$$

where i stands for industry and t for year.[6] Our identification relies on the fact that most of the variation in the NTR gap across tradeable industries arises from variation in non-NTR rates, set 70 years prior to passage of PNTR.[7] This feature mitigates concerns of reverse causality, that would arise if non-NTR rates could be set to protect struggling industries. Our difference-in-differences identification strategy exploits the large cross-sectional variation in the NTR gaps across U.S. tradeable industries in the years 1990–2001, *before* China was granted PNTR. We compare the stock returns of U.S. firms in high NTR gap industries to low NTR gap industries (first difference), during the uncertainty period, 1990–2001, versus the years 1980–1989 (second difference).[8] These potential tariff increases were substantial: in 1990 the average NTR gap across the tradeable industries in our sample was 27% with a standard deviation of 14%. The distribution of NTR gaps in 1990 is displayed in Fig. A.1 in Appendix A.

A priori, it is not obvious that the prospect of higher tariffs on Chinese goods was good news for U.S. tradeable industries. On one hand, if Congress actually voted to revoke China's NTR status, U.S. producers could have benefited in terms of reduced competition from Chinese firms. On the other hand, reverting to Smoot-Hawley tariffs could have hurt these industries if, for instance, firms sourced intermediate inputs from China, increasing production costs, or if China reciprocated with higher tariffs on U.S. goods, hurting U.S. exporters. In addition, investors did not know if/when the uncertainty would ever be resolved, i.e., if/when China would be granted PNTR. Therefore, U.S. firms were exposed also to the possibility of suddenly facing more competition from China. We argue that these considerations all generated uncertainty for U.S. firms, which we capture with the NTR gap.

To construct our sample of stock returns, we start with the universe of publicly listed U.S. firms in CRSP that can be matched to Compustat, which is where we obtain all the firm-level variables used as controls in the regressions. We then filter for ordinary common shares (share-codes 10 and 11) traded on major exchanges (NYSE, AMEX and NASDAQ). In order to have time-consistent industry definitions for tracking stock returns and other controls over our sample period, we use the

[4] U.S. president Jimmy Carter began granting such waiver to China annually in 1980, under the premises of the U.S. Trade Act of 1974.

[5] See Online Appendix of Pierce and Schott (2016) for several pieces of anecdotal evidence suggesting how the renewal of China's NTR status was perceived as uncertain.

[6] Pierce and Schott (2016) compute NTR gaps using ad-valorem equivalent NTR and non-NTR tariff rates from 1989 to 2001 provided by Feenstra et al. (2002). Both types of tariffs are set at the eight-digit Harmonized System (HS) level. Industry-level NTR gaps are then computed using concordances provided by the U.S. Bureau of Economic Analysis (BEA), such that the gap for an industry is the average NTR gap across the eight-digit HS tariff lines belonging to that industry.

[7] A regression of the NTR gap in 1990 on the non-NTR rate across industries gives a R^2 of 0.96, while a regression of the NTR gap in 1990 on the NTR rate in 1990 gives a R^2 of only 0.15.

[8] In a robustness exercise (see Table A.4), we also compare the uncertainty period to the post-PNTR period, with similar results.

algorithm developed in Pierce and Schott (2012) to create "families" of four-digit SIC industries. Unless otherwise noted, all references to "industry" in this paper refer to these families. As in Pierce and Schott (2016), we exclude all industries that have missing NTR gaps, i.e., non-tradeable industries. This also excludes industries that had positive NTR rates but missing non-NTR rates. We match the SIC code in Compustat to the Pierce and Schott (2012) families of industries and keep the matched firms. Table A.2 in Appendix A compares summary statistics for our final sample between high and low NTR gap firms.

2.2. Diff-in-Diff specification

We begin our analysis by estimating the following regression at the firm-month level:

$$r_{ity} = \alpha + \beta\ Uncertainty_t \times Gap_{i,y-1} + \delta_i + \delta_t + \boldsymbol{X}'_{it-1}\boldsymbol{\lambda} + \epsilon_{it} \tag{2}$$

where the dependent variable is the market-adjusted return (i.e., the return on the firm minus the return on the CRSP value-weighted index) of firm i in month t and year y, for the years 1980 to 2001. The first term on the right-hand side is the Difference-in-Differences (DID) term of interest, an interaction of the one year-lagged NTR gap of firm i and an indicator for the uncertainty period, i.e., equal 1 in the years characterized by tariff uncertainty, 1990–2001.[9] Therefore, the DID term of interest equals zero in the control period, 1980–1989. $\boldsymbol{X}_{it-1}$ is a vector of lagged industry-time controls, to be specified below, while δ_i and δ_t are firm and month fixed effects, which control for firm-specific unobserved characteristics and time trends in stock returns.

We use time-varying NTR gaps to prevent a look-ahead bias and to allow for time variation in the measure of uncertainty, in order to identify more precisely the response of stock returns over time.[10] We use the lagged NTR gaps to avoid endogeneity issues, which may arise if NTR tariff rates responded to contemporaneous changes in stock returns. In addition, the implicit assumption is that every year, investors' used previous-year NTR gaps to assess the level of each industry's tariff uncertainty. Within each month, observations are weighted by firms' stock market capitalization at the end of the previous month. This is consistent with the weighting scheme we use in our portfolio analysis, and prevents our estimates from being overly influenced by the performance of small firms. Standard errors are clustered at the firm level to allow for arbitrary error correlations within firms over time. The final sample consists of 5,655 firms, 281 tradeable industries over 252 months, for a total of 593,777 observations.

The baseline results are shown in Table 1. We report the coefficients in percentage points, and throughout the rest of the paper all returns are defined in percentage points, i.e., net returns multiplied by 100. Column (1) first considers a simple specification that includes only the DID term, the un-interacted NTR gap in 1990, and month fixed effects. We can see that the DID term is positive and 1% significant, suggesting that high-gap firms had higher average monthly returns than low-gap firms during the uncertainty period 1990–2001. The coefficient of 1.21 implies that a firm more exposed to trade policy uncertainty (i.e., a firm belonging to an industry at the 75th percentile of the distribution of NTR gaps in 1990) would have had yearly stock returns 3.19% higher than a firm less exposed to TPU (i.e., a firm belonging to an industry at the 25th percentile).

In column (2) we repeat the same specification but only include firms which we can match to our set of firm level controls: this is important to determine whether or not our results are driven by a selection effect, i.e., whether there is something systematically different about firms with non-missing financial data in Compustat. In column (3) we control for policy changes related to China's accession to the WTO that could have influenced the performance of U.S. industries over our sample period. To this end, we include the NTR tariff rates, Chinese import tariffs from Brandt et al. (2012), and data on U.S. textile and clothing quotas from Khandelwal et al. (2013). We also control for the industry exposure to the NAFTA shock, proxied by U.S. tariffs on Mexican goods in 1990 as in Hakobyan and McLaren (2016), interacted with year dummies.[11] In column (4) we add several firm-level financial characteristics known to be correlated with expected returns: the market beta (estimated in rolling 1-year windows), the natural logarithm of firm market capitalization, book to market, the P/E ratio, gross profitability and long-term leverage (see e.g., Fama and French (1996), Fama and French (2015)).[12] We can see that the coefficient on the DID term of interest remains positive and statistically significant throughout all the specifications, and although it is a bit lower, it is not statistically different from the coefficient in column (1). This suggests a limited role of the controls in affecting the magnitude of the results.

Lastly, in column (5) we also add firm fixed effects to account for unconditional differences in average returns across firms. The difference-in-differences coefficient of 2.58 is significant at 1% level, and it implies that increasing the exposure of a U.S. firm to TPU from the 25th to 75th percentile of the distribution of NTR gaps raises stock returns by 6.8% per year.

9 Although President Clinton signed the law granting PNTR in October 2000, China actually entered the WTO in December 2001, thus Congress voted also in 2001 on whether to revoke China's NTR rates.

10 Nevertheless, in Table A.4 we show that results are similar if we fix the NTR gap to its value in 1990.

11 Specifically, we use data on U.S. tariffs on imports from Mexico in 1990 collected by Feenstra et al. (2002), and aggregate the product-level tariffs at the industry level using U.S. imports from Mexico in 1990 as weights. Since all tariffs on Mexican goods went to 0 after NAFTA, the higher the average industry tariff, the stronger is the exposure to the NAFTA shock, because that meant a larger change in the exposure to Mexican imported goods.

12 The betas are obtained from the WRDS beta suite, while all the firm-level financial characteristics, except size, from the WRDS financial ratios suite. Firm market capitalization is computed using data from CRSP.

Table 1
TPU and stock returns.

Dep. variable:	Monthly returns, r_{it}				
	(1)	(2)	(3)	(4)	(5)
$Gap_{i,y-1} \times Uncertainty_t$	1.211***	1.090***	1.082***	1.072**	2.581***
	(0.318)	(0.324)	(0.369)	(0.447)	(0.614)
$Gap_{i,1990}$	−0.370	−0.370	−0.580	0.030	
	(0.370)	(0.380)	(0.456)	(0.521)	
R^2	0.011	0.012	0.012	0.014	0.036
Observations	593,777	522,406	522,406	522,406	522,400
Firm FE	N	N	N	N	Y
Month FE	Y	Y	Y	Y	Y
Controls	N	N	Y	Y	Y

Notes: This table contains selected estimates from versions of the following regression, run at the firm(i)/month(t) level:
$r_{it} = \alpha + \beta\ Uncertainty_t \times Gap_{i,y-1} + \delta_i + \delta_t + \boldsymbol{X}'_{it-1}\boldsymbol{\lambda} + \epsilon_{it}$
where $Uncertainty_t$ is a dummy equal to one if the year is between 1990 and 2001, r_{it} is the market-adjusted return of firm i in month t multiplied by 100. The regression also includes the following controls in $\boldsymbol{X}_{it-1}$: NTR tariff rates, Chinese import tariffs, quotas, exposure to NAFTA, market beta, estimated in one-year rolling windows, book-to-market, P/E ratio, gross profitability, long-term leverage, and firm size, measured by the natural logarithm of market capitalization. δ_i and δ_t are firm and month fixed effects. Within each month, observations are weighted by firm i's market capitalization in at the end of the previous month. Robust standard errors, clustered at the firm level, are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

This number is significantly higher than the effect implied by the specifications without firm fixed effects. This suggests that there is significant within-industry firm-level heterogeneity in returns that is not explained by the controls.[13] For this reason, we will refer to this specification as our baseline in the remainder of this Section.

2.3. Dynamics

For the differential stock performance of high-gap firms to be attributable to exposure to TPU, our policy measure, the NTR gap, should be positively correlated with stock returns only during the 1990–2001 period, but not in the control period. To determine whether there is a relationship between the NTR gap and stock returns in the years 1980–1989, we estimate a version of our baseline regression in rolling windows. We start by de-meaning returns both at the firm-level and at the month level (to account for the firm/month fixed effects in the baseline regression). We then regress these demeaned returns on the NTR gap and our control variables in 5-year rolling windows. The use of 5-year rolling windows is common in the finance literature, see e.g., Frazzini and Pedersen (2014), and it allows us to flexibly look at the dynamics of stock returns.

Fig. 1 plots the coefficients (where the coefficient at time t is estimated using data for the years $t-4$ to t) along with their 95% confidence intervals, clustered at the firm level. The graph shows that during the 1980–1989 decade there was no positive and significant association between the NTR gap and stock returns, suggesting the absence of any pre-trends. In contrast, we can see that the estimated coefficient for the first 5 years of the uncertainty period is positive, large and statistically different than zero. The coefficient stays large and positive for the following 4 years, then it drops and is no longer statistically different from zero. This pattern seems consistent with the fact that there was greater uncertainty about the legislation passage in the years 1990–1992, as suggested by the fact that in those years the House voted in favor of revoking NTR status to China (see Table A.1 in the Appendix). The fact that the coefficients becomes smaller and imprecise in the final years of the decade suggests that the investors perceived the policy uncertainty as fading away, as the likelihood of China joining the WTO was inevitably increasing over the years.[14]

2.4. Robustness

In Appendix A we perform several exercises to gauge the robustness of our baseline results. First, we follow Greenland et al. (2019) and construct an exposure to TPU at the firm level, by taking an average of the NTR gaps of the major business segments in which each firm operates, weighted by the firm' sales in each segment, obtained from Compustat Segment Data. Table A.3 estimates the same specifications as Table 1, but using this firm-level exposure to TPU. We can see that the coefficients are very close to the baseline ones.

Second, in Table A.4 we explore different specifications, using as baseline column (5) of Table 1. Column (1) reports the results of the baseline specification, but with standard errors clustered at both firm and month level, as in Petersen (2009), to

[13] To visualize this difference, Fig. A.2 in Appendix A plots the difference in average returns between the period of policy uncertainty (1990–2001) and the pre-period (1980–1989) at the industry-level (y-axis) vs. the firm-level (x-axis). Note that the figure simply compares differences in average residualized returns during the treated and non-treated periods, thus does not account for the effect of NTR gaps. If there was no within-industry variation in the change in average returns between the pre-period and the TPU period, all the points would lie on the 45 degree line. This, however, is not the case: while the relationship is upward sloping, and statistically significant, the slope is smaller than 1.

[14] In addition, the correlation between the real GDP per-capita growth (annual, averaged over 5 years) and the estimated diff-in-diff coefficients in Fig. 1 is −0.63. This suggests that the effect of TPU on stock returns is countercyclical, consistent with the predictions in Pástor and Veronesi (2013).

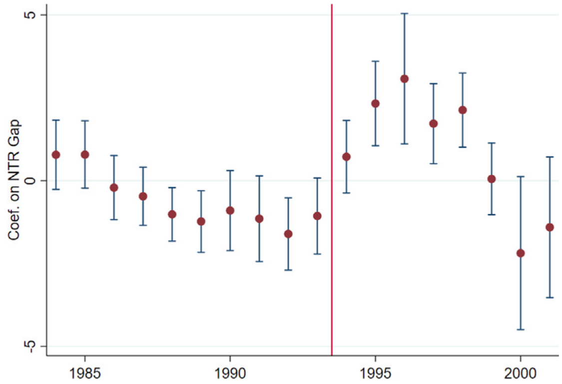

Fig. 1. 5-year rolling windows regressions. *Notes:* This figure is constructed in two steps: First, we de-mean returns both at the firm-level, and at the month level (to account for the firm/month fixed effects in the baseline regression). We then regress these demeaned returns on the NTR gap and our control variables in 5-year rolling windows. The blue lines surrounding the point estimates (red dots) are 95% confidence intervals, calculated based on standard errors clustered at the firm-level. Within each month, observations are weighted by a firm's lagged market capitalization.

account for the potential autocorrelation of the residuals. With double clustering, the standard error increases but the coefficient of interest is still significant at 1% level. Column (2) weighs observations with firms' market capitalization in December 1979, at the beginning of the period. This specification has far less observations than the baseline, because it requires a firm to be present in the sample in 1979, and implies a smaller DID coefficient, but still positive and highly significant. Column (3) includes an interaction between all control variables and the treatment period dummy, to ensure that the other covariates do not generate differential average returns during the uncertainty period (see e.g., Gagliardini et al. (2016)). The DID coefficient is very similar to the baseline.

One concern with our identification strategy is that the U.S. government could have set high NTR tariff rates to protect industries that they expected to perform poorly. In this case, while the non-NTR rates are exogenous because they were set by the Smoot-Hawley act, NTR rates are not. A second related concern is that the decision to vote on legislation to revoke China's temporary NTR status could have motivated by economic reasons, rather than geo-political reasons, i.e., the Tiananmen Square Massacre. This could generate an omitted variable problem in our regressions, leading to biased estimates. To mitigate these concerns, we follow Pierce and Schott (2016) and estimate a two-stage least squares specification in which we instrument the baseline DID term, $Uncertainty_t \times Gap_{i,y-1}$, with an interaction of the uncertainty indicator and the Smoot-Hawley non-NTR rates, $Uncertainty_t \times NonNTR_i$. As reported in column (4) of Table A.4, the DID coefficient is positive, statistically significant and close to the baseline.

Column (5) reports the results when holding the NTR gap at its level in 1990, the year the uncertainty started. Column (6) includes the post-PNTR period as control group, column (7) considers only firms present in all the years of the sample, column (8) controls for the loadings of the stocks on the TPU and EPU indices constructed in Baker et al. (2016).[15] In all these specifications, the coefficient on the NTR gap is 1% statistically significant and close to the baseline specification.

3. A risk premium for tariff uncertainty

Our empirical results show that high gap industries had higher average returns, relative to low gap industries, between 1990 and 2001. In this section, we argue that this difference in returns was a risk premium for exposure to tariff uncertainty. We first perform a portfolio analysis which shows that, *only* in the uncertainty period 1990–2001, a portfolio long high-gap sectors and short low gap industries had positive and significant returns. We then present additional empirical evidence on discount rate effects and realized volatility that support our risk premium hypothesis.

3.1. Conceptual framework

In any asset pricing model (see e.g., Cochrane (2009) and Duffie (2010)), the maximization problem of the representative investor implies the following relationship:

$$E(r_{it} - r_f) \propto -Cov(SDF_t, r_{it}) \tag{3}$$

Eq. (3) states that the expected excess return of any asset *i* is inversely proportional to the covariance of the asset's return with the Stochastic Discount Factor (SDF). Intuitively, stocks that covary negatively with the SDF are riskier because they pay

[15] Additional results not reported for brevity show that our findings are robust if i) we use the industry classification in CRSP, and ii) we exclude the computer and electronics industries that experienced the dot-com crash in 2000–2001, iii) we extend the sample to 2017.

less in bad states of the world, and thus require higher expected returns as compensation for risk-averse investors.[16] Empirically, expected returns are estimated using average returns, computed over long periods of time.

To fully characterize the SDF, further assumptions about the economy are needed. The key elements in defining the SDF are the factors that affect the marginal utility of consumption. Given our setting, we think it is appropriate to start with the well-known frameworks in Pastor and Veronesi (2012) and Pástor and Veronesi (2013). In these models, stock dividends depend on the government policy, which is uncertain, and thus any policy change affects the future marginal utility. This implies that government policy enters the SDF, i.e., investors require an additional compensation for exposure to policy uncertainty. In the context of our analysis, this means that stocks more exposed to TPU, i.e., stocks with higher NTR gap, should have a higher (in absolute value) covariance with the component of the SDF associated with TPU. From Eq. (3), this implies higher expected returns.[17]

In the next section, we follow the vast literature on empirical asset pricing (see e.g., Nagel (2013) and Fama and French (2015)) and test this hypothesis by ranking stocks according to their exposure to TPU and forming value-weighted portfolios. We then present additional empirical evidence, consistent with the predictions of the Pastor and Veronesi (2012) model, that supports our risk premium hypothesis.

3.2. Portfolio analysis

In order to estimate the risk premium associated with exposure to TPU, we first rank the firms in our sample in 3 subgroups, based on their associated NTR gap in the previous year. We construct value-weighted portfolios, where each month, the weights are proportional to each firm's market capitalization in the previous month. Finally, we calculate monthly returns for these portfolios between 1990 and 2001. We also construct a "Trade Policy Uncertainty" (TPU) portfolio, which is the difference in returns between the portfolios containing firms with the highest and lowest gaps, divided by two. We then estimate the following regression, separately for each portfolio p:[18]

$$r_t^p = \alpha^p + \boldsymbol{F}_t'\boldsymbol{\beta} + \epsilon_t \tag{4}$$

where r_t^p is the excess return on portfolio p in month t, and $\boldsymbol{F}_t$ is a vector containing the 5 Fama and French (2015) factors: the market portfolio minus the risk-free rate, the size factor (small minus big), the value factor (high minus low), the profitability factor (robust minus weak), and the investment factor (conservative minus aggressive).[19] Our hypothesis is that firms in industries more exposed to policy uncertainty should command higher expected returns, because investors require compensation for such exposure, as in Pastor and Veronesi (2012). If our hypothesis is correct, then the estimated constant α^p in Eq. (4) should be: i) monotonically increasing as we go from low to high gap portfolios, and ii) positive and significant for the TPU portfolio.

Our results in Table 2 show that the TPU portfolio had an α^p of 0.515 percentage points per month during the uncertainty period, significant at the 1% level. Furthermore, the constant is monotonically increasing as we go from the low gap to high gap portfolios. The estimated coefficient implies that the TPU portfolio, long high-gap firms and short low-gap firms, would have earned 6.18% per year throughout the 1990–2001 period. Therefore, trade policy uncertainty was a systematic factor that could *not* be diversified away across stocks with similar NTR gaps.

3.2.1. Robustness and extensions

We examine the robustness of our results in Table A.5 in Appendix A. For brevity we only report the results for the TPU portfolio (long high-gap and short low-gap). Column (1) reports the results obtained by sorting the portfolios with the firm-level NTR gaps constructed using the Compustat Segments data. Column (2) reports the results using the baseline industry-level gaps but restricting to firms that are present every month between 1990 and 2001. In both specifications, the estimated alpha is lower than the baseline one, although not statistically different from it. We think that this is partly due a selection effect that arises from using the Compustat Segments data. As evidenced in Table A.4, there are many firm-year observations in Compustat for which we do not have information about their segments. On top of that, sometimes firms fail to report some of their operating segments (see Cohen and Lou (2012)). Finally, a segment only has to be reported if it comprises at least 10% of a firm's consolidated annual sales, which means the segments data will not capture all sales. Columns (3) and (4) repeat the baseline analysis but using the 1980–1989 and 2001–2007 periods. During the control periods, high-gap industries did not have economically large or statistically significant alphas, suggesting that the risk premium was earned only during the uncertainty period, consistent with our hypothesis. Lastly, column (5) constructs factor-mimicking

[16] For example, with CRRA utility and time-separable preferences, the SDF is proportional to the marginal utility of consumption. See also footNote 21.

[17] This conclusion implicitly assumes that: i) the covariance between the component of the SDF associated with jumps in TPU and stock returns is negative, i.e., increases in TPU are perceived as bad states of the world by investors, and ii) the loading on the TPU factor in the SDF is positive. Both endogenously hold in Pastor and Veronesi (2012) (see Proposition 6), and are consistent with a vast literature arguing that increases in uncertainty are detrimental for the economy (see e.g., Bloom et al. (2007), Bloom (2009) and Bachmann et al. (2013)).

[18] The implicit assumption is that the loadings on other risk-factors, such as the Fama and French (2015) factors, are randomly distributed across high and low gap firms. In addition, under standard assumptions on the stochastic process for the returns, one can show that the SDF is linear in its arguments (see e.g., Pastor and Veronesi (2012)). This implies that by constructing a portfolio that goes long high-gap firms, and short low-gap firms, we can effectively isolate exposure to TPU.

[19] We obtain the monthly returns on these factors from Ken French's website: https://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html.

Table 2
Portfolio analysis.

Dep. variable:	Monthly returns, r_t^p			
	Low Gap	Medium Gap	High Gap	TPU
Market	0.901***	0.980***	0.964***	0.032
	(0.044)	(0.057)	(0.064)	(0.045)
Size	−0.054	0.122*	0.114	0.084*
	(0.052)	(0.073)	(0.072)	(0.047)
Value	−0.015	−0.310***	−0.385***	−0.185**
	(0.085)	(0.104)	(0.114)	(0.084)
Profitability	0.231***	0.06	−0.185**	−0.208***
	(0.064)	(0.084)	(0.090)	(0.060)
Investment	0.472***	−0.066	−0.421**	−0.447***
	(0.108)	(0.141)	(0.164)	(0.116)
α^p	−0.374**	0.249	0.656***	0.515***
	(0.146)	(0.198)	(0.205)	(0.141)
Observations	144	144	144	144
R^2	0.80	0.84	0.89	0.71
Sample Period	1990–2001	1990–2001	1990–2001	1990–2001

Notes: This table contains selected estimates from the following regression, using data from 1990–2001: $r_t^p = \boldsymbol{F}_t'\boldsymbol{\beta} + \alpha^p + \epsilon_t$.
where r_t^p is the return on value-weighted portfolio p in month t. $\boldsymbol{F}_t$ is a vector containing the 5 Fama–French factors: Market, Size, Value, Profitability and Investment. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

portfolios for exposure to the economic policy uncertainty and trade policy uncertainty series in Baker et al. (2016). To do this, we regress monthly returns on changes in the Baker et al. (2016) series and keep the betas. We then rank firms into 3 groups based on the betas, and construct portfolios which go long the high beta firms and short the low beta firms. Our results survive controlling for these factor-mimicking portfolios, suggesting that the trade policy uncertainty captured by the NTR gap is orthogonal to the text-based measures of Baker et al. (2016) and Brogaard and Detzel (2015). This is further confirmed in Fig. A.3, which shows that there is no significant relationship between the NTR gap and a firms' beta on the TPU factor from Baker et al. (2016).

A concern with our results is that high-gap industries may have had higher returns than low gap industries because of unobserved industry-level shocks that occurred between 1990 and 2001. If this were the case, however, we would expect to see high gap industries outperforming low gap industries in other countries, despite a lack of exposure to trade policy uncertainty. To rule out this possibility, we repeat the baseline portfolio analysis using monthly industry-level stock returns for 6 high-income countries: Japan, South Korea, the United Kingdom, France, Germany and Australia.[20] Table 3 reports that a portfolio which goes long high-gap firms, and goes short low-gap firms, i.e., our TPU portfolio, did not have a significantly positive average returns in any of the 6 high-income countries considered between 1990 and 2001. This suggests the difference in stock returns between high and low gap industries was specific to the United States, corroborating our risk-premium explanation.[21]

Finally, we investigate how exposure to TPU interacts with exposure to globalization. We first look at the effect of trade protection, forming portfolios by double-sorting the firms according to both the NTR gap in 1990 and the NTR tariff rate in 1990. To do this, we calculate the median of the NTR gap in 1990, and the median of the NTR rate in 1990. We then do a 2-by-2 sort to form 4 total portfolios. The first two columns of Table 4 show that the risk premium was earned mostly by firms with a low level of NTR rate, i.e., firms operating in industries less protected by trade policy from foreign competition. In a similar vein, in the next two columns we measure exposure to globalization with the industry-level shipping costs used in Barrot et al. (2018). When we double sort portfolios on the NTR gap in 1990 and on the shipping costs from Barrot et al. (2018), we find that the TPU risk premium was earned by industries with lower shipping costs, i.e., sectors that were exposed more to globalization. These findings suggest that lack of protection from import competition, either in the form of higher tariffs or higher shipping costs, can significantly amplify the risk premium for exposure to trade policy uncertainty.

Lastly, we study another channel through which TPU may have affected the riskiness of U.S. firms. The last two columns of Table 4 document that, within high-gap industries, industries with a higher share of inputs expenditures from China throughout the 1990–2001 period earned a larger risk premium.[22] This suggests that uncertainty about the cost of production,

[20] We only include countries for which we have a good coverage of the time series and cross-sectional dimensions in the Compustat Global database.

[21] Additional evidence in favor of the risk premium hypothesis comes from the CCAPM. Under power utility, the risk premium is simply proportional to $-Cov\left(\beta(C_{t+1}/C_t)^{-\gamma}, r_{it}\right)$, where γ is the relative risk aversion of the consumer, r_{it} is the return of asset i, and C_{t+1}/C_t is the growth rate of aggregate consumption. We estimate this correlation with a range of relative risk aversion from 1 to 30, gross growth of U.S. private consumption expenditure, the S&P500 market portfolio, and our TPU portfolio. We find that the correlation between the SDF and the TPU portfolio is between −0.20 and −0.22, and statistically significant between 1990 and 2001. Instead, in the 1980–1989 and 2001–2007 control periods such correlation is not statistically significant.

[22] We use data from WIOD to compute the share of expenditures of each downstream U.S. industry on each upstream industry from China. Since the sectors in the WIOD are more aggregated than the industries in our sample, we assume that this share is constant across industries within each WIOD sector.

Table 3
Portfolio analysis, foreign countries.

Dep. variable:	Monthly returns, r_t^p					
	TPU	TPU	TPU	TPU	TPU	TPU
Market	0.145***	0.296***	0.082	0.082*	−0.345***	0.251**
	(0.039)	(0.104)	(0.052)	(0.044)	(0.082)	(0.115)
α^p	0.001	1.069	−0.503	0.016	0.847	−0.943**
	(0.261)	(0.857)	(0.370)	(0.360)	(0.627)	(0.468)
Observations	144	144	144	144	144	144
R^2	0.077	0.086	0.016	0.031	0.179	0.031
Country	Japan	South Korea	UK	France	Germany	Australia

Notes: This table contains selected estimates from the following regression: $r_t^p = \boldsymbol{F}_t'\boldsymbol{\beta} + \alpha^p + \epsilon_t$.
where r_t^p is the return on the value-weighted TPU portfolio in month *t*. $\boldsymbol{F}_t$ contains a country-specific market factor, which is the value-weighted average return of all the firms in Compustat Global for that country. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 4
TPU factor and exposure to globalization.

Dep. variable:	Monthly returns, r_t^p					
	Low NTR TPU	High NTR TPU	Low Ship TPU	High Ship TPU	Low Inputs TPU	High Inputs TPU
Market	0.014	−0.008	0.016	0.093**	0.041	0.032
	(0.058)	(0.038)	(0.085)	(0.042)	(0.065)	(0.064)
Size	0.139**	0.045	0.145	0.090*	0.034	0.03
	(0.063)	(0.042)	(0.095)	(0.050)	(0.078)	(0.060)
Value	−0.016	−0.179**	0.394**	−0.282***	0.101	−0.046
	(0.100)	(0.088)	(0.158)	(0.073)	(0.098)	(0.102)
Profitability	0.064	−0.026	−0.472***	0.179**	−0.313***	−0.082
	(0.062)	(0.049)	(0.117)	(0.071)	(0.089)	(0.069)
Investment	−0.635***	0.165*	−0.973***	0.141	−0.345*	−0.599***
	(0.158)	(0.098)	(0.204)	(0.108)	(0.180)	(0.150)
α^p	0.585***	0.061	0.528**	0.017	0.012	0.465**
	(0.192)	(0.125)	(0.253)	(0.124)	(0.182)	(0.183)
Observations	144	144	144	144	144	144
R^2	0.47	0.12	0.47	0.266	0.372	0.54

Notes: This table contains selected estimates from the following regression, using data from 1990–2001: $r_t^p = \boldsymbol{F}_t'\beta + \alpha^p + \epsilon_t$.
where r_t^p is the return on the value-weighted TPU portfolio in month *t*. $\boldsymbol{F}_t$ is a vector containing the 5 Fama–French factors: Market, Size, Value, Profitability and Investment. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

deriving from uncertainty on the tariffs imposed on intermediate inputs from China, is a risk factor that was also priced by financial markets.

3.3. Discount rate effect

In Section 3.1, we argue that high-gap firms had higher average returns than low-gap firms between 1990 and 2001 because investors required a risk premium for exposure to trade policy uncertainty. In any asset pricing model, the stock price of firm *i* in period *t* equals:

$$P_{i,t} = \frac{D_{i,t+1}}{r_i - g_i}, \tag{5}$$

where $D_{i,t+1}$ is the cash-flow at $t+1$, r_i is a weighted-average of all future discount rates, and g_i is a weighted-average of all future growth rates (see e.g., Campbell and Shiller (1988)). The discount rate embeds firm *i*'s loadings on all priced risk factors which, as argued in Section 3.1, should include trade policy uncertainty. Eq. (5) implies that if there is a sudden jump in

[23] The Tiananmen Square Massacre happened 3 days after President Bush's MFN extension recommendation for 1989, and no resolution was introduced in the Congress in 1989 to disapprove the President's recommendation. The U.S. response to Tiananmen Square that year consisted of two sets of non-MFN-related sanctions announced by President Bush (on June 5 and June 20, 1989), and a series of bills that Congress considered which were designed to codify the President's actions and expand their scope. See Dumbaugh (1998).

[24] See https://www.congress.gov/bill/101st-congress/house-bill/4939.

[25] Of course, another reason for the drop in stock price could have been a relative decrease in the next period's dividends, $D_{i,t+1}$. However, the literature has shown that uncertainty typically has long-run effects that would not show up instantaneously on dividends (see e.g., Bloom (2009) and Barrero et al. (2017)). In fact, when we look at dividends paid in 1991, they actually increased relatively more for high-gap firms.

TPU, the discount rate increases and the stock price decreases. Thus, if we could identify the date when the uncertainty on China's NTR status began, we should observe a sudden decrease in stock prices for exposed industries. As shown in Pastor and Veronesi (2012), this discount rate effect at the announcement of a new policy should be stronger for industries more exposed to TPU, i.e., the ones with higher NTR gap.

We look at the stock price responses around the days in which the uncertainty about China's tariff status reasonably began. In particular, while the Tiananmen Square Massacre happened on June 4, 1989, the first resolution to revoke China's NTR status was introduced in the House on May 24, 1990, by Rep. Donald Pease (H.R. 4939), and it was reported by the Committee on Ways and Means on July 23, 1990 (H. Rept. 101–620).[23] As introduced, the bill would have directed the President to take new conditions – involving substantial progress on human rights violations – into account when extending China's MFN status beginning in 1991 (see Dumbaugh (1998)).[24]

We examine value-weighted portfolios of high-gap and low-gap firms, defined as having NTR gaps above/below the median in 1990. We work with market-adjusted returns of these portfolios to take out the common mean component over the period of interest. Fig. 2 documents the value of $1 invested in each of these portfolios on May 24, 1990. We can see that, from when the bill was introduced until July, 1990, high and low gap firms were following the same trend. On July 18, 1990 the bill was ordered to be reported by the Committee on Ways and Means, and it was actually reported by the Committee on July 23, 1990 (H. Rept. 101–620). Since then, i.e., when it became known to the public that the House was considering revoking China's NTR status, high-gap firms' market-adjusted stock returns started to decrease relative to low-gap firms. This suggests a strong discount rate effect pushing down prices for tradeable sectors more exposed to the newly introduced trade policy uncertainty. Interestingly, when the bill was actually approved by the House on October 18, 1990, there was no significant impact on stock prices, suggesting that the markets had already priced in the increased TPU.

Another explanation for the drop in stock prices for high NTR gap firms could have been a decrease in g_i.[25] This may have occurred, for example, if investors believed that increased policy uncertainty would have led high-gap firms to hold cash, instead making productive investments, relatively more than low-gap firms. If investors had perfect foresight, the effect of increased uncertainty on all future expected dividend growth rates should have been reflected in prices. To test this hypothesis, we calculate realized dividend growth rates for the 3 portfolios formed in Section 3.2. We find that between 1990 and 2001, the dividend growth rate of high-gap firms was not statistically significantly different from the one of low-gap firms. Similarly, we find no relationship between net income growth between 1990 and 2001 and the NTR Gap, as shown in Fig. A.4 in the Appendix. This is suggestive evidence that high gap firms did not have particularly strong realized financial performance during the period of trade policy uncertainty. Therefore, even with perfect foresight, a change in g_i was likely not responsible for the price drop in Fig. 2.

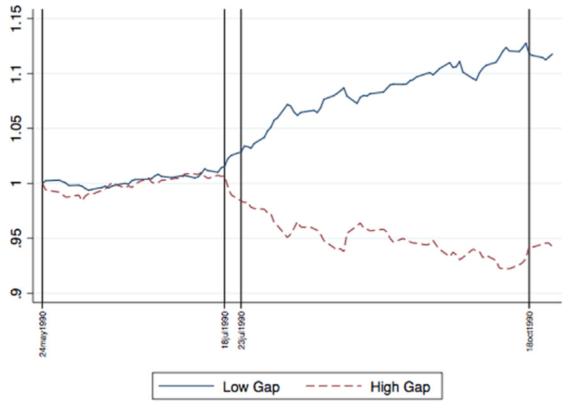

Fig. 2. Effect of Introducing Uncertainty. *Notes:* The figure plots the value of 1 dollar invested in value-weighted portfolios of high and low gap industries on 5/24/1990.

[23] The Tiananmen Square Massacre happened 3 days after President Bush's MFN extension recommendation for 1989, and no resolution was introduced in the Congress in 1989 to disapprove the President's recommendation. The U.S. response to Tiananmen Square that year consisted of two sets of non-MFN-related sanctions announced by President Bush (on June 5 and June 20, 1989), and a series of bills that Congress considered which were designed to codify the President's actions and expand their scope. See Dumbaugh (1998).

[24] See https://www.congress.gov/bill/101st-congress/house-bill/4939.

[25] Of course, another reason for the drop in stock price could have been a relative decrease in the next period's dividends, $D_{i,t+1}$. However, the literature has shown that uncertainty typically has long-run effects that would not show up instantaneously on dividends (see e.g., Bloom (2009) and Barrero et al. (2017)). In fact, when we look at dividends paid in 1991, they actually increased relatively more for high-gap firms.

3.4. Realized volatility

We investigate whether uncertainty about China's trade status was also associated with more volatile stock prices around House voting days. Intuitively, regardless of whether the policy change is perceived as good or bad by investors, firms more exposed to such policy changes should have larger responses (in absolute terms) than less exposed firms, as argued both theoretically and empirically in Pastor and Veronesi (2012), Boutchkova et al. (2012) and Baker et al. (2019).

We test this hypothesis with the following regression:

$$|r_{it}| = \alpha + \theta_1\ Gap_{it} + \theta_2\ D_t + \theta_3\ D_t\ HighGap_{it} + \epsilon_{it} \tag{6}$$

where $|r_{it}|$ is the absolute market-adjusted return of firm i in day t, winsorized at the 1% and 99% level by year. We use the absolute return, instead of the squared return (the quadratic variation), as our preferred measure of realized volatility to prevent a small number of outliers from driving our results. D_t is a vector of three dummy variables equal to 1 if t is: (1) 3 days before the event date (2) on the event date itself and (3) 3 days after the event date. Gap_{it} is the NTR gap and $HighGap_{it}$ is a dummy variable equal to one if the firm had an above-median NTR gap. Within each day, observations are weighted by lagged market capitalization. We focus on all days in which Congress voted to revoke the NTR status to China, from 1990 to 2001.

Column (1) in Table 5 documents that high gap firms had higher than average volatility, and in particular were relatively more volatile both before, on and after voting dates. The higher volatility of more exposed firms *before* the voting days may be associated with "political uncertainty", i.e., investors were uncertain about which way the vote will go, and it is consistent with the predictions of Pastor and Veronesi (2012). The higher volatility of high-gap firms on and after the voting days may be associated with "policy uncertainty", i.e., given the way the vote went, investors were uncertain about the implications for high-gap firms. Again, this is consistent with the predictions of Pastor and Veronesi (2012).

In order to ensure that the higher unconditional volatility of high-gap firms does not drive the result on the event dates, we select the median dates between the true voting dates, as placebo announcement days each year. Column (2) shows that the interaction term on these placebo days is not significant, confirming that the higher volatility of high-gap firms is specific to key House voting days.

4. Alternative explanations

In this section we discuss a number of alternative explanations that could potentially rationalize our results, and show that they are not consistent with the empirical evidence. This additional set of results provides strong evidence in favor of our explanation that exposure to TPU is a source of risk that is priced in the cross-section of stock returns.

4.1. Expected cash-flow effect

A potential explanation for our results is that the higher returns for high-gap industries in the uncertainty period were driven by the response of stock prices to news and policy-related events that affected the expected cash-flows of the firms, rather than by a change in the risk premium. To disentangle the effect of expected cash flows from the pure uncertainty effect, we adopt two different empirical strategies, and both lend support to our hypothesis.

Table 5
Realized volatility around voting days.

Dep. variable:	Realized Volatility	
	Voting days	Placebo days
$Gap_{i,y-1}$	2.135***	2.153***
	(0.020)	(0.020)
$Pre \times Gap_{i,1990}$	0.116***	0.009
	(0.037)	(0.035)
$Day \times Gap_{i,1990}$	0.240***	−0.001
	(0.070)	(0.048)
$Post \times Gap_{i,1990}$	0.152***	−0.025
	(0.047)	(0.036)
R^2	0.044	0.044
Observations	8,340,591	8,340,591

Notes: This table contains selected estimates from versions of the following regression, run at the firm(i)/day(t) level using data from 1990–2001: $|r_{it}| = \alpha + \theta_1\ Gap_{it} + \theta_2\ D_t + \theta_3\ D_t\ HighGap_{it} + \epsilon_{it}$.
where $|r_{it}|$ is the absolute market-adjusted return for firm i on day t. D_t is a vector of three dummy variables equal to 1 if t is: (1) 3 days before the event date (2) on the event date itself and (3) 3 days after the event date. $Gap_{i,t}$ is the NTR gap and $HighGap_{i,t}$ is a dummy variable equal to one if the firm had an above-median NTR gap. Within each day, observations are weighted by lagged market capitalization. Robust standard errors, clustered at the firm level, are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

First, we repeat our portfolio analysis but exclude a window of 3 days before and after each policy-related event. If capital markets are efficient, stock prices adjust quickly after a news announcement, incorporating any changes in expected future cash-flows, as recently shown by Breinlich (2014) and Greenland et al. (2019). We focus on all days in which Congress voted to revoke the NTR status to China, from 1990 to 2001, and on two key policy announcements: i) 10/10/2000, when China was granted permanent NTR, conditional on joining the WTO; ii) 12/11/2001, when China joined the WTO. We sum daily log-returns to compute monthly returns.

Table 6 documents that excluding days around key policy dates does not significantly alter the estimated risk premium. The point estimate is lower than the baseline, but not statistically different from it, and it implies a risk premium of 3.58%. One reason the risk premium is lower than the baseline is that we compound the daily returns using logs, which by construction smooths out large variation in returns. In fact, when we repeat the baseline portfolio analysis, without excluding any policy event, but using log returns, we find a risk premium of about 3% per year. Therefore, we fail to find evidence that stock prices responses around policy days affect our estimates.

Another way to separate out the expected cash-flow effects is to explicitly look at firms that were hurt the most after the uncertainty was resolved. To this end, we compute, at the firm level, the difference in EBIT growth between 1990–2001 and 2002–2007. We focus on EBIT because it is a measure insensitive to the capital structure of the firm. Lower values mean that EBIT grew slower (or was shrinking) in the post-PNTR period, relative to the 1990's. We interact this variable with the NTR gap and a dummy variable for the period of policy uncertainty. If the investors knew that high gap firms would have been economically hurt by the resolution of uncertainty (as shown ex-post by Pierce and Schott (2016)), the interaction term should be negative and significant.

Table 7 shows, instead, that there is no difference in average returns between firms that ex-post did poorly and ex-post did well after PNTR went into effect. We also find that including this interaction term does not qualitatively change our coefficient of interest, relative to Table 1.[26]

4.2. Realized returns at earnings announcements

A potential concern with our results is that the differences in stock returns we pick up across high and low gap industries are driven by returns around quarterly earnings announcements, reflecting investors' forecast errors on firms' performance after the policy change in 1990. To mitigate this concern, we follow Barrot et al. (2018) and test whether the differential returns of our high and low gap portfolios were concentrated around earnings announcements. We identify earnings days using the Institutional Brokers Estimates System (I/B/E/S) database.[27] We look at the cumulative returns from $t - 3$ to $t + 3$, where t is an earnings announcement date.

We repeat the baseline portfolio analysis shown in Section 3.2, but only including the days of firms' earnings announcements. Table 8 documents that the difference in returns around earnings announcement days for stocks in the low and high gap portfolios is not statistically significant.[28] These findings provide evidence against the hypothesis that the ex-post average realized returns that we measure in the data deviate in a systematic way from the unobserved ex-ante expected returns for investors.

4.3. Protection from China

One possible explanation for the baseline results is that high-gap industries were receiving higher de facto protectionism from China via the uncertainty generated by the trade policy status. There is in fact evidence that these industries suffered in terms of employment and investment relatively more when the uncertainty was eliminated in 2001 (see Pierce and Schott (2016), Handley and Limão (2017) and Pierce and Schott (2018)). This would have translated into higher expected cash flows for high-gap sectors, and thus in higher average returns between 1990 and 2001. We have already shown in Table 7 that there is no significant relationship between the growth rate of cash-flows after 2001 and firms stock returns. Here we focus specifically on the effect of Chinese competition.

To disentangle the effect of the implicit protection from China given by the uncertainty from the effect of TPU per se on stock returns, we examine whether the effects of TPU were larger or smaller in industries with low Chinese import penetration, i.e., industries more "protected" by policy uncertainty. We measure import penetration as the industry-year level import share from China.[29] If this alternative hypothesis is correct, we should observe that: i) average returns are higher in industries that were more protected from China, i.e., industries that had a lower import share from China throughout 1990–

[26] Note that this exercise has fewer observations and different point estimates from Table 1 because it requires a firm to have at least two years of non-missing EBIT data in Compustat in the post-2001 period, in order to compute the change in average growth rates.

[27] If earnings are announced after the market is closed, or on a trading holiday, we set the effective earnings day to the first trading-day after earnings are announced.

[28] These results also also confirmed by a simple regression of stock returns around earnings days in the period 1990–2001 on the NTR gap in 1990. We find an insignificant effect of more exposure to TPU on stock returns around earning days.

[29] We use U.S. import data at the HS-6 level downloaded from the Center for International Data at UC Davis, https://cid.econ.ucdavis.edu/usix.html. We aggregate the data at the industry level, and then compute the imports from China as share of total production, obtained from the NBER Manufacturing database. Results are similar if we divide the imports from China by the total imports of the industry.

Table 6
Portfolio analysis excluding policy days.

Dep. variable:	Monthly returns, r_t^p			
	Low Gap	Medium Gap	High Gap	TPU
Market	0.894***	0.969***	0.959***	0.032
	(0.049)	(0.060)	(0.063)	(0.043)
Size	−0.04	0.13	0.114	0.077*
	(0.062)	(0.082)	(0.083)	(0.046)
Value	−0.024	−0.323***	−0.394***	−0.185**
	(0.101)	(0.123)	(0.127)	(0.080)
Profitability	0.230***	0.093	−0.152	−0.191***
	(0.077)	(0.095)	(0.137)	(0.071)
Investment	0.478***	−0.101	−0.394**	−0.436***
	(0.129)	(0.172)	(0.187)	(0.109)
α^p	−0.649***	−0.382*	−0.052	0.298**
	(0.161)	(0.212)	(0.204)	(0.132)
Observations	144.000	144.000	144.000	144.000
R^2	0.77	0.82	0.86	0.70
Sample Period	1990–2001	1990–2001	1990–2001	1990–2001

Notes: This table contains selected estimates from the following regression, using data from 1990–2001: $r_t^p = \boldsymbol{F}'_t\beta + \alpha^p + \epsilon_t$.
where r_t^p is the return on value-weighted portfolio p in month t. $\boldsymbol{F}_t$ is a vector containing the 5 Fama–French factors: Market, Size, Value, Profitability and Investment. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 7
Stock Returns and EBIT growth.

Dep. variable:	Monthly returns, r_{it}				
	(1)	(2)	(3)	(4)	(5)
$Gap_{i,y-1} \times Uncertainty_t$	1.303***	1.153***	1.089***	1.099**	2.119***
	(0.355)	(0.369)	(0.415)	(0.489)	(0.618)
$Gap_{i,1990}$	−0.285	−0.242	−0.319	0.153	
	(0.395)	(0.423)	(0.496)	(0.557)	
Diff in EBIT growth x NTR Gap	−0.058	−0.021	−0.015	−0.161	0.108
	(0.282)	(0.308)	(0.309)	(0.291)	(0.634)
R^2	0.017	0.018	0.018	0.020	0.039
Observations	345,812	304,172	304,172	304,172	304,169
Firm FE	N	N	N	N	Y
Month FE	Y	Y	Y	Y	Y
Controls	N	N	Y	Y	Y

Notes: This table contains selected estimates from versions of the following regression, run at the firm(i)/month(t) level: $r_{it} = \alpha + \beta\, Uncertainty_t \times Gap_{i,y-1} + \delta_i + \delta_t + \boldsymbol{X}'_{it-1}\boldsymbol{\lambda} + \epsilon_{it}$.
where $Uncertainty_t$ is a dummy equal to one if the year is between 1990 and 2001, r_{it} is the return of firm i in month t. The regression also includes the following controls in $\boldsymbol{X}'_{it}$: NTR tariff rates, Chinese import tariffs, quotas, exposure to NAFTA, market beta, estimated in one-year rolling windows, book-to-market, P/E ratio, gross profitability, long-term leverage, and firm size, measured by the natural logarithm of market capitalization. δ_i and δ_t are firm and month fixed effects. Within each month, observations are weighted by firm i's market capitalization in at the end of the previous month. Robust standard errors, clustered at the firm level, are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

2001, and ii) the risk premium, i.e., the return of the high-minus-low portfolio, was concentrated in industries with lower import share from China.

To assess this hypothesis, we first sort firms into two groups: those in industries with above/below the median import share from China. Then, within these groups, we form two sub-groups: those in industries with above/below median NTR gap. Finally, we construct a portfolio which goes long high-gap industries and short low-gap industries i.e., a TPU portfolio among those industries with low import share from China, and a TPU portfolio among those industries with high import share from China.

Table 9 documents two facts. First, α^p is negative for low-gap firms, irrespective of the level of imports from China (columns (1) and (3)), while α^p is positive for high-gap firms, irrespective of the level of imports (columns (3) and (4)). Second, there is a non-significant α^p for industries with less import penetration from China. Instead, the risk premium was concentrated in industries *less* protected from China, as indicated by the positive and 1% significant α^p of the TPU portfolio formed among high imports sectors, i.e., industries that were importing relatively more from China. This evidence is also consistent with the previous findings in Table 4, which shows that the risk premium was earned mostly by industries more vulnerable

Table 8
TPU factor around earnings announcements.

Dep. variable:	Monthly returns, r_t^p			
	Low Gap	Medium Gap	High Gap	TPU
Market	0.269***	0.320***	0.237***	−0.016
	(0.050)	(0.075)	(0.080)	(0.040)
Size	0.089**	−0.005	0.145**	0.028
	(0.036)	(0.075)	(0.067)	(0.030)
Value	0.063	−0.074	−0.044	−0.054
	(0.056)	(0.112)	(0.133)	(0.063)
Profitability	0.072	−0.046	−0.055	−0.063
	(0.049)	(0.085)	(0.086)	(0.045)
Investment	0.064	−0.165	−0.157	−0.111
	(0.090)	(0.138)	(0.187)	(0.087)
α^p	−0.343***	−0.523**	−0.041	0.151
	(0.112)	(0.201)	(0.203)	(0.102)
Observations	144.000	144.000	144.000	144.000
R^2	0.37	0.45	0.35	0.21
Sample Period	1990–2001	1990–2001	1990–2001	1990–2001

Notes: This table contains selected estimates from the following regression, using data from 1990–2001: $r_t^p = \boldsymbol{F}_t'\beta + \alpha^p + \epsilon_t$. where r_t^p is the return on value-weighted portfolio p in month t. $\boldsymbol{F}_t$ is a vector containing the 5 Fama–French factors: Market, Size, Value, Profitability and Investment. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 9
TPU factor and imports from China.

Dep. variable:	Monthly returns, r_t^p					
	Low gap		High gap		Low Imp.	High Imp.
	Low Imp.	High Imp.	Low Imp.	High Imp.	TPU	TPU
Market	0.915***	1.108***	1.059***	0.973***	0.072	−0.067
	(0.057)	(0.058)	(0.075)	(0.101)	(0.053)	(0.055)
Size	−0.117**	0.212***	0.083	0.058	0.101	−0.077
	(0.058)	(0.068)	(0.098)	(0.102)	(0.069)	(0.064)
Value	−0.393***	0.045	−0.390***	0.061	0.002	0.008
	(0.101)	(0.114)	(0.142)	(0.147)	(0.100)	(0.101)
Profitability	0.349***	0.105	−0.096	−0.237**	−0.222**	−0.171***
	(0.078)	(0.071)	(0.152)	(0.091)	(0.094)	(0.057)
Investment	0.839***	0.157	0.069	−0.914***	−0.385**	−0.536***
	(0.167)	(0.161)	(0.223)	(0.223)	(0.159)	(0.150)
α^p	−0.107	−0.436**	0.405	0.665**	0.256	0.550***
	(0.187)	(0.168)	(0.267)	(0.271)	(0.171)	(0.159)
Observations	144	144	144	144	144	144
R^2	0.724	0.844	0.778	0.843	0.479	0.43
Sample Period	1990–2001	1990–2001	1990–2001	1990–2001	1990–2001	1990–2001

Notes: This table contains selected estimates from the following regression, using data from 1990–2001: $r_t^p = \boldsymbol{F}_t'\beta + \alpha^p + \epsilon_t$. where r_t^p is the return on value-weighted portfolio p in month t. $\boldsymbol{F}_t$ is a vector containing the 5 Fama–French factors: Market, Size, Value, Profitability and Investment. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

to globalization. Therefore, this evidence suggests that the higher returns of high-gap sectors we observe in the 1990–2001 period were not due to more implicit protection from China.[30]

[30] In our reasoning, we are assuming that the implicit protection from China given by TPU translates into lower import shares. Of course, differences in import penetration across sectors may be due to other factors as well, such as comparative advantage. However, suppose that the level of import penetration in 1990–2001 depended only on comparative advantage. Then, we should observe that U.S. industries with higher imports from China had lower stock returns, because of lower expected dividends. Instead, Table 9 documents the opposite, i.e., *higher* average returns. In addition, note that imports from China after 2001 increased relatively more in industries with lower initial import penetration, as documented in Fig. A.5 in the Appendix. The coefficient of the relationship is significant at 1% level.

5. Conclusions

We use quasi-experimental variation arising from China's temporary NTR status to show that U.S. tradeable industries more exposed to trade policy uncertainty had significantly higher stock returns than less exposed industries between 1990 and 2001. Our measure of uncertainty, which relies on the difference between current NTR and non-NTR tariff rates, has the advantage of being directly observable, quasi-exogenous, and ex-ante. As such, it is not subject to the concerns associated with ex-post measures of uncertainty typically used in the finance literature.

Our estimated risk premium is substantial, even after accounting for the response of stock prices at key policy-related dates, and for the effect of Chinese import competition on stock returns. Industries highly exposed to policy uncertainty earned a risk premium of 3.6–6.2% per year relative to less exposed sectors, suggesting a large impact of trade policy uncertainty on the perceived riskiness of exposed stocks. Among industries more exposed to trade policy uncertainty, the risk premium was larger if shipping costs were low, the NTR tariff rate was low, or if the industry relied more heavily on imports from China. These amplification channels make intuitive sense: uncertainty about trade policy matters more if an industry is directly exposed to globalization, either through low trade barriers or through the supply chain.

Our focus in this paper is on trade policy uncertainty before China entered the WTO, but currently the U.S.-China trade relationships are also unstable. The large risk premia should have informed policy makers at the time, and today. Our approach is general: policy makers and practitioners can use risk premia to understand the effect of protracted uncertainty on investor's beliefs and stock market behavior.

CRediT authorship contribution statement

Marcelo Bianconi: Conceptualization, Methodology, Software, Investigation, Formal analysis, Writing – original draft, Writing – review & editing. **Federico Esposito:** Conceptualization, Methodology, Software, Validation, Formal analysis, Investigation, Data curation, Writing – original draft, Writing – review & editing, Visualization. **Marco Sammon:** Conceptualization, Methodology, Software, Validation, Formal analysis, Investigation, Data curation, Writing – original draft, Writing – review & editing, Visualization.

Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Appendix A

See Figs. A.1–A.5 and Tables A.1–A.5.

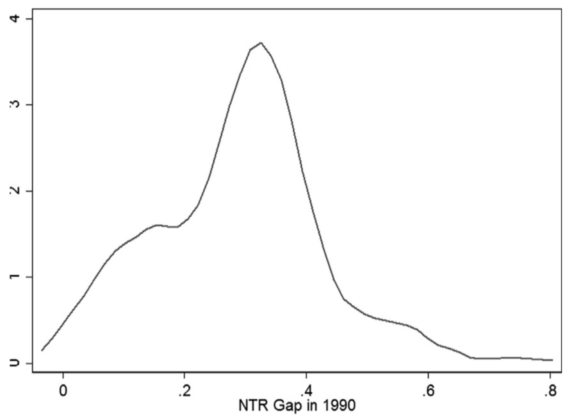

Fig. A.1. Distribution of NTR gaps in 1990. *Notes:* The figure displays the empirical distribution of the NTR gaps in 1990.

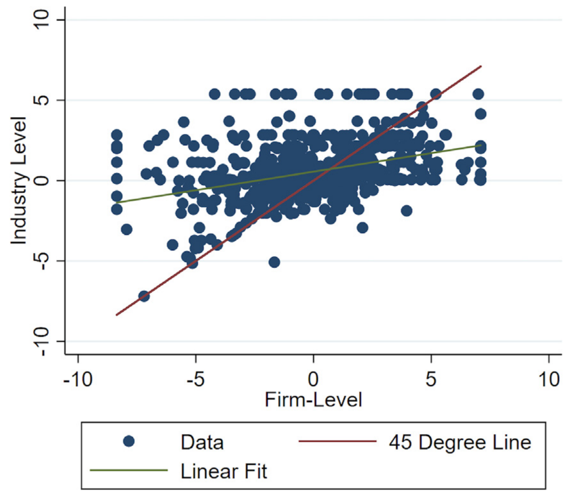

Fig. A.2. Differences in average returns, 1980–1989 vs 1990–2001. *Notes:* The figure plots the difference in average residualized returns, in percentage points, between the period of policy uncertainty (1990–2001) and the pre-period (1980–1989) at the industry-level (y-axis) vs. the firm-level (x-axis). Both quantities are winsorized at the 1.

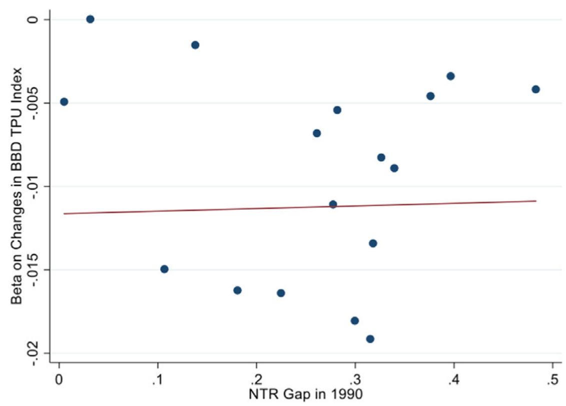

Fig. A.3. NTR gaps and Baker et al. TPU. *Notes:* The figure displays a binned scatter plot where the y-axis is the coefficient from a regression of a firm's monthly returns on monthly changes in the Baker et al. 2016 trade policy uncertainty index, estimated between 1990 and 2001. The x-axis is the firm's NTR gap in 1990.

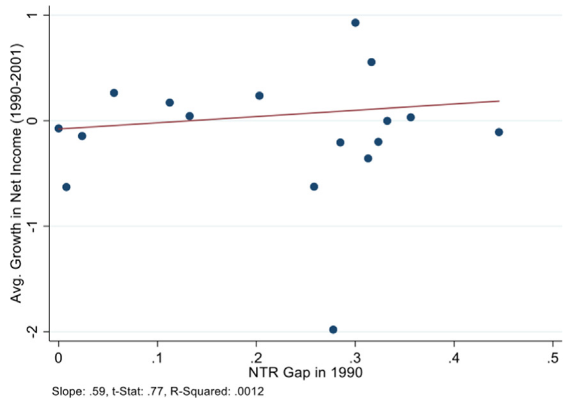

Fig. A.4. NTR gaps and income growth. *Notes:* The figure displays a binned scatter plot where the y-axis is the average growth in net income (obtained from Compustat balance sheet data for each firm) between 1990 and 2001. The x-axis is the firm's NTR gap in 1990.

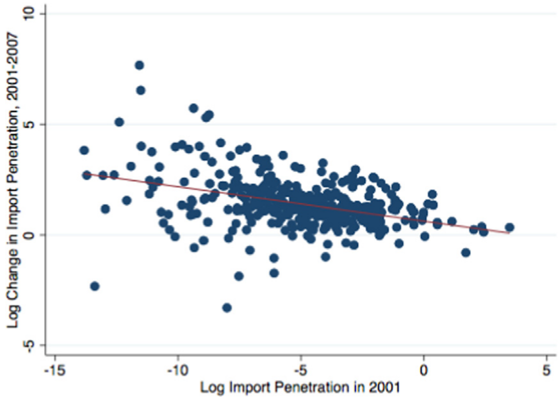

Fig. A.5. Change in import penetration from China. *Notes:* The figure plots the log-change in the industry-level import penetration from China between 2011 and 2007, on the log of the import penetration in 2001. Import penetration is computed as imports from China as share of total production. Results are similar if we divide the imports from China by the total imports of the industry.

Table A.1
Outcome of house votes to revoke NTR.

	Disapprove	Approve	%Disapprove
1990	247	174	57
1991	223	204	51
1992	258	135	59
1993	105	318	24
1994	75	356	17
1995	107	321	26
1996	141	286	32
1997	173	259	40
1998	166	264	38
1999	170	260	39
2000	147	281	34
2001	169	259	39
mean	165	260	38

This table displays annual votes in the U.S. House of Representatives to renew China's NTR status. Total number of possible votes is 435. Source: Pierce and Schott (2016).

Table A.2
Summary statistics.

Variable	Low-Gap	High-Gap	t-statistic
NTR Gap in 1990	0.14	0.31	−37.97
Market beta	0.50	0.52	−0.87
Book to Market	0.78	0.80	−0.66
P/E	11.00	9.35	0.99
Profitability	0.30	0.40	−6.45
Leverage	0.85	0.75	0.76
Log of size	10.94	10.58	3.81

Notes: This table contains summary statistics on high and low gap firms in 1990. A firm is classified as low-gap if it has a below median NTR gap in 1990. Each entry represents the un-weighted average within each group. The last column contains the t-Statistic from a difference of means test across groups.

Table A.3
TPU and stock returns, segments data.

Dep. variable:	Monthly returns, r_{it}				
	(1)	(2)	(3)	(4)	(5)
$Gap_{i,y-1} \times Uncertainty_t$	1.26***	1.24***	1.19***	0.97**	2.75***
	(0.310)	(0.319)	(0.340)	(0.419)	(0.714)
$Gap_{i,1990}$	0.110	0.140	0.060	0.590	
	(0.321)	(0.323)	(0.340)	(0.429)	
R^2	0.010	0.011	0.011	0.012	0.027
Observations	547,440	489,761	489,761	489,761	489,755
Firm FE	N	N	N	N	Y
Month FE	Y	Y	Y	Y	Y
Controls	N	N	Y	Y	Y

Notes: This table contains selected estimates from versions of the following regression, run at the firm(i)/month(t) level: $r_{it} = \alpha + \beta\ Uncertainty_t \times Gap_{i,y-1} + \delta_i + \delta_t + \boldsymbol{X}'_{it-1}\boldsymbol{\lambda} + \epsilon_{it}$.
where $Uncertainty_t$ is a dummy equal to one if the year is between 1990 and 2001, r_{it} is the return of firm i in month t. The regression also includes the following controls in $\boldsymbol{X}_{it}$: NTR tariff rates, Chinese import tariffs, quotas, exposure to NAFTA, market beta, estimated in one-year rolling windows, book-to-market, P/E ratio, gross profitability, long-term leverage, and firm size, measured by the natural logarithm of market capitalization. δ_i and δ_t are firm and month fixed effects. Within each month, observations are weighted by firm i's market capitalization in the previos month. Robust standard errors, clustered at the firm level, are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.4
TPU and stock returns, robustness.

Dep. variable:	Monthly returns, r_{it}							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$Gap_{i,y-1} \times Uncertainty_t$	2.58***	1.32**	2.47***	2.50***	1.74**	2.98***	2.37***	2.54***
	(0.699)	(0.631)	(0.730)	(0.610)	(0.816)	(0.530)	(0.606)	(0.640)
$BBD_1 \times Uncertainty_t$								49.71***
								(11.230)
$BBD_2 \times Uncertainty_t$								3.47
								(3.392)
R^2	0.036	0.081	0.037	0.012	0.035	0.031	0.092	0.036
Observations	522,400	215,432	522,400	522,400	522,705	655,366	78,248	510,992
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y
Month FE	Y	Y	Y	Y	Y	Y	Y	Y
Controls	Y	Y	Y	Y	Y	Y	Y	Y

Notes: Column (1) clusters errors at firm and month level, column (2) weights observations with market capitalization in 1979, column (3) adds an interaction between the controls and the uncertainty dummy, column (4) uses an IV methodology, column (5) uses the NTR Gap in 1990, column (6) includes the 2002–2007 peiod, column (7) includeds only firms present in all the years of the sample, column (8) controls for the loading of stock on the TPU index from Baker et al. 2016 (BBD_1) and on the EPU index (BBD_2). Within each month, observations are weighted by firm i's market capitalization in the previous month, except in column (2). Robust standard errors, clustered at the firm level, are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.5
Portfolio analysis, robustness.

Dep. variable:	Monthly returns, TPU portfolio				
	(1)	(2)	(3)	(4)	(5)
Market	0.03	0.042	−0.009	0.062	0.009
	(0.039)	(0.048)	(0.028)	(0.076)	(0.042)
Size	0.015	0.004	0.170***	0.087	−0.031
	(0.047)	(0.052)	(0.060)	(0.077)	(0.047)
Value	−0.338***	−0.202**	−0.288***	−0.222*	−0.342***
	(0.086)	(0.095)	(0.080)	(0.124)	(0.086)
Profitability	0.021	−0.027	−0.255***	−0.328**	0.097
	(0.049)	(0.072)	(0.086)	(0.135)	(0.071)
Investment	−0.141	−0.212*	−0.025	0.144	−0.106
	(0.094)	(0.118)	(0.118)	(0.142)	(0.095)
BBD EPU					−0.008
					(0.068)
BBD TPU					−0.128
					(0.082)
α^p	0.348***	0.366**	0.187	0.017	0.328***
	(0.116)	(0.140)	(0.144)	(0.155)	(0.121)
Observations	144	144	120	72	144
R^2	0.56	0.39	0.39	0.51	0.57
Sample Period	1990–2001	1990–2001	1980–1989	2002–2007	1990–2001

Notes: This table contains selected estimates from the following regression: $r_t^p = \boldsymbol{F}_t'\beta + \alpha^p + \epsilon_t$.
where r_t^p is the return on the value-weighted TPU portfolio in month t. $\boldsymbol{F}_t$ is a vector containing the 5 Fama–French factors: Market, Size, Value, Profitability and Investment. Robust standard errors are in parenthesis. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

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