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Three Essays in Macro-Finance, International Economics and Macro-Econometrics

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À ma tante Pauline, ma maman Martine et mon feu papa Lévis

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Résumé

Cette thèse présente de nouveaux résultats sur différentes branches de la litérature en macro-finance, économie internationale et macroéconométrie. Les deux premiers chapitres combinent des modèles théoriques et des techniques empiriques pour approfondir l'étude de phénomènes économiques importants tels que les effets de l'incertitude liée aux politiques économiques sur les marchés financiers, et la convergence entre les pays émergents et les pays avancés sur ces marchés. Le troisième chapitre, qui est le fruit d'une collaboration avec Hafedh Bouakez, contribue à la litérature sur l'identification des chocs anticipés sur la productivité future.

Dans le premier Chapitre, j'étudie l'effet de l'incertitude relative aux politiques monétaire et fiscale sur les rendements et les primes de risque associés aux actifs nominaux du gouvernement des États-Unis. J'utilise un modèle d'équilbre stochastique et dynamique de type néo-Keynesien prenant en compte des préférences récursives des agents et des rigidités réelles et nominales. En utilisant un modèle VAR structurel. L'incertitude relaticve aux poliques économiques est définie comme étant une expansion de la distribution des chocs de politique, expansion au cours de laquelle la moyenne de la distribution reste inchangée. Mes résultats montrent que: (i) Lorsque l'économie est sujette à des chocs imprévisibles sur la volatilité des instruments de politique, le niveau médian de la courbe des rendements baisse de 8,56 points de base, sa pente s'accroît de 13,5 points de base et les primes de risque baissent en moyenne de 0.21 point de base. Cet effet négatif sur le niveau des rendements et les primes de risque est dû à l'impact asymétrique des chocs de signes opposés mais de même amplitude; (ii) Un choc positif à la volatilité des politiques économiques entraîne une hausse des rendements pour toutes les durées de maturité. Cet effet s'explique par le comportement des ménages qui, à la suite du choc, augmentent leur demande de bons dans le but de se prémunir contre les fortes fluctuations espérées au niveau de la consommation, ce qui entraîne des pressions à la baisse sur les rendements. De façon simultanée, ces ménages requièrent une hausse des taux d'intérêt en raison d'une espérance d'inflation future plus grande. Les analyses montrent que le premier effet est dominant, entraînant donc la hausse des rendements observée. Enfin, j'utilise plusieurs mesures empiriques d'incertitude de politiques économiques and un modèle VAR structurel pour montrer les résultats ci-dessus sont conformes avec les faits empiriques.

Le Chapitre 2 explore le marché des bons du gouvernement de 12 pays avancés et 8 pays

émergents, pendant la période 1999-2012, et analyse la question de savoir s'il y a eu une quelconque convergence du risque associé à ces actifs entre les deux catégories de pays. Je fais une distinction entre risque de défaut et autres types de risque, comme ceux liés au risque d'inflation, de liquidité ou de change. Je commence par montrer théoriquement que le différentiel au niveau des primes de risque "forward" entre deux pays peut être utilisé pour faire la distinction entre le risque de défaut et les autres risques. Ensuite, je construis ces différentiels de risque "forward" et les utilise pour montrer qu'il est difficile de conclure que ces autres types de risque dans les pays émergents ont convergé vers les niveaux qu'on observe dans les pays développés. Je montre aussi que la différence entre les fondamentaux macroéconomiques de ces deux catégories de pays, de même que des niveaux différents de risque politique, jouent un rôle important dans l'explication des différences de primes de risque—autres que celles associées au risque de défaut—entre les pays émergents et les pays avancés.

Le Chapitre 3 propose une nouvelle stratégie d'identification des chocs technologiques anticipés et non-anticipés, qui conduit à des résultats similaires aux prédictions des modèles néo-Keynésiens conventionnels. Il montre que l'incapacité de plusieurs méthodes empiriques à générer des résultats rejoignant la théorie est due à l'impureté des données existences sur la productivité totale des facteurs (TFP), conduisant à une mauvaise identification des chocs technologiques non-anticipés—dont les effects estimés ne concordent pas avec l'interprétation de tels chocs comme des chocs d'offre. Ce problème, à son tour, contamine l'identification des chocs technologiques anticipés. Mon co-auteur, Hafedh Bouakez, et moi proposons une stratégie d'identification agnostique qui permet à la TFP d'être affectée de façon comtemporaine par deux chocs surprises (technologique et non-technologique), le premier étant identifié en faisant recours aux restrictions de signe sur la réponse de l'inflation. Les résultats montrent que les effets des chocs technologiques anticipés et non-anticipés concordent avec les prédictions des modèles néo-Keynésiens standards. En particulier, le puzzle rencontré dans les travaux précédents concernant les effects d'un choc non-anticipé sur l'inflation disparaît lorsque notre nouvelle stratégie est employée.

Mots-clés: Incertitude de politique économique, Courbe des rendements, Volatilité stochastique, Préférences récursives, VAR structurel, Primes de risque, Marchés Émergents, Convergence, Risque Politique, Identification, Chocs technologiques anticipés, Restrictions de signes, Technologie, Productivité totale des facteurs.

Abstract

This thesis brings new evidence on different strands of the literature in macro-finance, international economics and macroeconometrics. The first two chapters combine both theoretical models and empirical techniques to deepen the analysis of important economic phenomena such as the effects of economic policy uncertainty on financial markets, and convergence between emerging market economies and advanced economies on these markets. The third chapter of the thesis, which is co-authored with Hafedh Bouakez, contributes to the literature on the identification of news shocks about future productivity.

In the first chapter, I study the effect of monetary and fiscal policy uncertainty on nominal U.S. government bond yields and premiums. I use a New-Keynesian Dynamic Stochastic General Equilibrium model featuring recursive preferences, and both real and nominal rigidities. Policy uncertainty in the DSGE model is defined as a mean-preserving spread of the policy shock distributions. My results show that: (*i*) When the economy is subject to unpredictable shocks to the volatility of policy instruments, the level of the median yield curve is lower, its slope increases and risk premiums decrease relative to an economy with no stochastic volatility. This negative effect on the level of yields and premiums is due to the asymmetric impact of positive versus negative shocks; (*ii*) A typical policy risk shock increases yields at all maturities. This is because the fall in yields triggered by higher demand for bonds by households, in order to hedge against higher predicted consumption volatility, is outweighed by the increase in yields due to higher inflation risk premiums. Finally, I use several empirical measures economic policy uncertainty in a structural VAR model to show that the above effects of policy risk shocks on yields are consistent empirical evidence.

Chapter 2 looks at the market for government bonds in 12 advanced economies and 8 emerging market economies, during the period 1999-2012, and consider the question of whether or not there has been any convergence of risk between emerging market and advanced economies. I distinguish between default risk and other types of risk, such as inflation, liquidity and exchange rate risk. I make the theoretical case that forward risk premium differentials can be used to distinguish default risk and other risks. I then construct forward risk premium differentials and use these to make the empirical case that there has been little convergence associated with the other types of risk. I also show that

differences in countries' macroeconomic fundamentals and political risk play an important role in explaining the large "non-default" risk differentials observed between emerging and advanced economies.

Chapter 3 proposes a novel strategy to identify anticipated and unanticipated technology shocks, which leads to results that are consistent with the predictions of conventional new-Keynesian models. It shows that the failure of many empirical studies to generate consistent responses to these shocks is due to impurities in the available TFP series, which lead to an incorrect identification of unanticipated technology shocks—whose estimated effects are inconsistent with the interpretation of these disturbances as supply shocks. This, in turn, contaminates the identification of news shocks. My co-author, Hafedh Bouakez, and I propose an agnostic identification strategy that allows TFP to be affected by both technological and non-technological shocks, and identifies unanticipated technology shocks via sign restrictions on the response of inflation. The results show that the effects of both surprise TFP shocks and news shocks are generally consistent with the predictions of standard new-Keynesian models. In particular, the inflation puzzle documented in previous studies vanishes under the novel empirical strategy.

Keywords : Economic policy uncertainty, Yield curve, Stochastic volatility, Recursive preferences, Structural VAR, Risk premiums, Emerging Markets, convergence, Political risk, Identification, News shocks, Sign restrictions, Technology, Total factor productivity.

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Chapter 1

What are the Effects of Economic Policy Uncertainty on Yields and Premia?

1.1 Introduction

I use a New-Keynesian DSGE model similar to that of Fernández-Villaverde et al. (2011), Born & Pfeifer (2014) and Mumtaz & Zanetti (2013) to answer the following two questions: (1) *How do yields and premiums on government bonds behave when the economy constantly faces unpredictable shocks to the volatility of policy instruments?* and (2) *What is the response of bond yields to a surprise policy risk shock?*. The paper, thus, contributes to bridge the gap between the flourishing literature on the macroeconomic effects of economic policy uncertainty on the U.S. economy, and the small number of studies that focus on its effects on yields and premiums. I also provide empirical evidence supporting the results of my theoretical model, using a structural vector autoregressive (VAR) model. By economic policy uncertainty I mean economic agents' expectations that monetary and fiscal authorities may, in the future, deviate from their standard behavior and take extreme measures in order to face special economic circumstances such as crises. I pay particular attention to the effects of monetary and fiscal policy uncertainty on yields and premiums. The role of recursive preferences and investment adjustment costs is also emphasized.

The motivation for this paper is that economists are paying more attention to the effect of policy risk, but most papers are focused on their real business cycle implications (see for example Fernández-Villaverde et al. (2011), Born & Pfeifer (2014), Mumtaz & Zanetti (2013), Leduc & Liu (2012))[1]. Yet, recent events, both political and economic, have shown that monetary and fiscal policy uncertainty may also have considerable implications for interest rate dynamics. Rudebusch et al. (2006) claimed that reduced uncertainty about U.S. monetary policy might have lowered the risk of holding long-term bonds. This could explain the fact that long-term yields fell in 2004-2005 despite a 150 basis points rise in short-term interest rates during the same period. This decoupling of both ends of the U.S. yield curve was referred to as a conundrum by the former chairman of the Federal Reserve, Alan Greenspan. This reduced uncertainty refers to the fact that the FED had become more transparent about the conduct of monetary policy during the previous two decades. Policy uncertainty may therefore influence the way investors perceive interest rate risk. Another example of the effects that policy uncertainty may have on the yield curve is related to the several debt ceiling episodes in the U.S. Interest rates on treasuries usually surge during these periods and one of the usual explanations, apart from the risk of default, is increased uncertainty about fiscal policy[2].

Studying the implications of policy uncertainty is of high relevance. Studying its effects on yields is important for policymakers as they are interested in understanding how changes in interest rates (due to uncertainty and other shocks) affect the real economy

[1] Throughout the paper, the expressions "policy risk" and "policy uncertainty" are used interchangeably.

[2] The Economic Policy Uncertainty index of Baker et al. (2013) rises considerably during these periods. The index of fiscal uncertainty constructed in Section 1.5 also shows higher values and increased volatility since the outbreak of the financial and economic crises and during the different debt ceiling episodes.

and public finances. The "Greenspan conundrum" mentioned above is a very good example of why the implications of policy uncertainty for bond prices is a relevant concern for policymakers, especially central bankers. This is important because, according to standard New-Keynesian models widely used by central banks, long-term rates are the channel through which monetary policy transmits to the real economy (see Rotemberg & Woodford (1999)).

The present paper studies the implications of a wide range of sources of policy uncertainty for nominal yields and premiums on zero-coupon bonds. My New-Keynesian DSGE model includes fiscal rules for each of the following fiscal variables: government spending, consumption tax rate, labor income tax rate and capital income tax rate, each of which features a stochastic volatility process. It also includes a Taylor-type monetary policy rule with a similar volatility process. Stochastic volatility of fiscal and monetary policy shocks helps account for a salient fact uncovered by many empirical studies using U.S. time series: time-varying standard deviations of innovations to fiscal and monetary policy rules. To mention a few of them, Primiceri (2005), Sims & Zha (2006), Born & Pfeifer (2014) and Mumtaz & Zanetti (2013) provide empirical evidence of these time variations, which is inconsistent with common practice in the majority of empirical and theoretical studies. Investment adjustment costs are shown to play an important role in driving medium and long yields' response to policy risk shocks. Price adjustment costs and the central bank's stabilization mechanism through a standard Taylor Rule drive inflation expectations, which are one of the key determinants of the response of yields and premiums to the shocks. Recursive Epstein & Zin (1989) preferences also help get the shape of the yield curve right. Rudebusch & Swanson (2012) show that a DSGE model in which agents have Epstein-Zin preferences, by helping the model produce large and variable risk premiums while also fitting key macroeconomic variables, can help solve a common puzzle according to which standard DSGE models display risk premiums that are too small and stable relative to their empirical counterparts [3]. Moreover, Bansal & Yaron (2004) find that including recursive preferences and risks related to varying growth prospects and fluctuating economic uncertainty can help the model replicate many features of asset prices, such as large and variable risk premiums.

Only a few papers study the implications of policy uncertainty for interest rates. From a modelling perspective, many authors represent a policy risk shock as a mean-preserving spread of the distribution of future policy instruments[4]. Koeda & Kato (2010) construct a no-arbitrage GARCH affine term structure model in which the volatility of the monetary policy shock is heteroskedastic, and find that an increase in monetary policy uncertainty raises the medium- and longer-term spreads. Mumtaz & Zanetti (2013) use a DSGE model, enriched with stochastic volatility of monetary policy, to provide a theoretical underpinning to their empirical findings that short-term interest rates, output growth

[3] This is known in the literature as the *risk premium puzzle* (see Backus et al. (1989), den Haan (1995)).

[4] Policy uncertainty can also be modelled by including time-variation to the reaction coefficient of policy rules as Buraschi et al. (2014) do. In this paper I do not focus on these alternative ways of modelling policy uncertainty

and inflation fall in reaction to an increase in the volatility of monetary policy. According to their results, higher monetary policy uncertainty increases the rate at which bonds gain value over time, leading economic agents to require lower compensations in order to hold bonds. Jordà & Salyer (2003) use a limited participation model to show that greater uncertainty about monetary policy induces agents to increase liquidity in the banking sector, leading to a decrease in short-term rates; longer-term rates also decrease due to the effect that the greater uncertainty has on the nominal intertemporal rate of substitution. The most closely related paper to mine is Leippold & Matthys (2015) as it explores the response of the yield curve to both government and monetary policy uncertainties. These authors estimate an affine yield curve model to study the impact of economic policy uncertainty on the term structure of nominal interest rates. They find that higher government policy uncertainty leads to a decline in yields due to the adverse effect that it has on the trend component of real output growth, that in turn renders capital investment more risky and induces investors to favor government bonds.

Relative to the literature cited above, my contribution is twofold. First, I show that the response of yields to policy risk shocks depends on the assumptions about investment adjustment costs. Indeed, most of the papers in the literature ignore this feature. Bloom (2009) stresses that ignoring these adjustment costs leads to substantial bias in the effect of uncertainty shocks. Moreover, Eberly et al. (2012) show that theoretical models featuring the type of investment adjustment costs that I use replicates a salient property (the lagged-investment effect) of firm level investment data in the U.S. My results highlight the fact that when agents face costs in adjusting investment, medium- and long-maturity bond yields increase following a policy risk shock. I show that this is because following a policy uncertainty shock, agents face a need for resource reallocation which is very difficult if there are investment adjustment costs. I show that uncertainty about capital tax policies is dominant, and these costs prevent agents from adjusting to the distortions that such a policy uncertainty creates, by reducing investment and increasing demand for bonds. In contrast, absent these investment adjustment costs, medium and long bond yields decrease because households decrease investment easily, and increase demand for bonds. I also show that adjustment costs lead to the fiscal effect dominating the monetary effect whereas absent these costs, the fiscal effect is dampened and eventually muted.

My second major contribution is that I put a particular emphasis on fiscal policy uncertainty which proves to have at least as much importance for yields dynamics as monetary policy uncertainty that has been the focus in the literature. This is also relevant because as many economists have emphasized—especially based on the various debt ceiling episodes and the fiscal cliff in the U.S.—uncertainty about fiscal policy can have sizable economic effects. My paper, thus, goes beyond analyzing the impact of monetary policy uncertainty as Mumtaz & Zanetti (2013) and Jordà & Salyer (2003) do, and includes fiscal policy uncertainty in a way that captures actual government spending and taxation decisions (unlike Leippold & Matthys (2015)). Indeed, I model the government's behavior based on solid micro-foundations and empirical evidence. The model specifies policy rules for government spending and tax rates; it uses stochastic volatility to the innovations of these instruments, which is consistent with empirical studies (see Born & Pfeifer (2014), and Fernández-Villaverde et al. (2011)). In contrast, Leippold & Matthys (2015) include an ad-hoc process that supposedly captures government's policy uncertainty and that directly affects productivity and real output growth. In practice, what they call a government policy uncertainty shock cannot be distinguished from a shock to, say, human capital or innovation.

While Leippold & Matthys (2015) study the same questions as the ones tackled in this paper, the two model setups are very different, leading to completely different results. The main departures of my model from theirs is the explicit modelling of the government behavior, and the inclusion of investment adjustment costs. The two features completely switch the sign of yield responses to policy risk shocks. On the one hand the introduction of an actual government leads to a switch in the sign of the response of short yields which become positive, due to a very high inflation risk premiums charged by investors on short-term bonds following a policy risk shock. On the other hand, the introduction of investment adjustment costs in the model reveals an additional transmission channel (smoothing channel) that Leippold & Matthys (2015)'s model fails to uncover. My empirical model results are in line with my theoretical findings, showing that the channels uncovered in this paper do play very important roles.

I calibrate the model to the U.S. economy using parameters mostly estimated in the related literature using quarterly data from 1970 to 2010. I solve the model using a third-order perturbation around the deterministic steady state and use it to answer the two questions asked above. Simulations from the model show that, compared to an environment which does not feature policy uncertainty, an uncertain economic policy environment triggers a decline of the level of the median yield curve by 8.56 basis points, increases its slope by 13.54 basis points and decreases term premiums on average by 0.21 basis points. These declines can be explained by the asymmetric effects of positive versus negative policy risk shocks, negative shocks (perceived as higher policy transparency) having larger effects than positive shocks. Impulse response analysis shows that all yields increase in response to a positive policy risk shock, but the higher the maturity the lower the initial impact on yields. The intuition behind this finding is that, due to investment adjustment costs, the fall in yields triggered by higher bond demand from households in order to hedge against higher predicted consumption volatility (the smoothing channel), is outweighed by the increase in yields due to higher inflation risk premiums (the inflation channel). As a consequence investors require higher yields to hold bonds. I show that reducing agents' investment adjustment costs improves their ability to transfer wealth intertemporally, emphasizing the smoothing channel. Absent these adjustment costs, the smoothing channel becomes dominant in the medium- and long-run and the corresponding yields decrease following the shock.

I then use a structural VAR model to investigate empirically the effects of monetary and fiscal policy uncertainty on yields. One of the challenges in doing this is to find

an appropriate measures of policy uncertainty. I use Forecasters disagreement on future values of policy-related variables such as the Consumer Price Index (CPI) and the short-term interest rate—for monetary policy—or government purchase of goods and services and the budget balance—for fiscal policy. I measure disagreement between forecasters by taking cross-sectional coefficients of variation between their one-year forecasts of the above variables, using data from the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters. Using the constructed indexes of monetary and fiscal policy uncertainty in a structural VAR, I identify policy risk shocks using sign restrictions à la Mountford & Uhlig (2009); in order to be consistent with a wide range of both empirical and theoretical findings, I impose that output responds negatively to a positive (monetary or fiscal) policy uncertainty shock. Consistently with the results of my theoretical model, the impulse response functions from the empirical model show that policy uncertainty shocks leads an increase in interest rates of all maturities. These results are robust to alternative measures of policy uncertainty (including Baker et al. (2013)'s measures), different model specifications, sample periods and lag length in the reduced-form VAR.
The remainder of this paper is organized as follows. Section 3.2 presents the model economy. In section 3.3, I show how nominal yields and term premia are derived, I also present the calibration of the model and present the solution method. The results of our investigations are presented and discussed in section 2.5, the transmission mechanism of policy uncertainty to yields and the important role of investment adjustment costs are also presented in the same section. Section 1.5 presents empirical evidence that supports the findings of the theoretical model. Section 2.6 concludes and provides some policy implications.

1.2 The Model Economy

The economy is populated by a representative household that works n_t hours, consumes c_t and invests in physical capital K_t and in a variety of nominal government bonds of different maturities, $\{B_t^l\}_{l=1}^L$. There is also a productive sector comprising a continuum of monopolistically competitive firms that produce intermediate goods by hiring labor and buying capital services from the household, and an aggregate good firm that uses intermediate goods to produce a composite good. Finally there is a government that levies taxes on consumption expenses, on labor and capital incomes, and issues nominal bonds in order to finance its spending.

1.2.1 The Household

The representative household maximizes its utility represented recursively and given by:

$$V_t = \left[(1-\beta)\left(c_t^{\nu}\left(1-n_t\right)^{1-\nu}\right)^{\frac{1-\gamma}{\theta}} + \beta\left(E_t V_{t+1}^{1-\gamma}\right)^{\frac{1}{\theta}}\right]^{\frac{\theta}{1-\gamma}}. \tag{1.1}$$

Where E_t is the expectation conditional on time t information set, β is the discount factor, ν is used to control labor supply, γ is the risk aversion parameter, and $\theta \equiv \frac{1-\gamma}{1-\frac{1}{\psi}}$ where ψ is the intertemporal elasticity of substitution (IES). In this framework, recursive preferences prove crucial for getting the shape of the yield curve right [5].

The resources of the representative household are: lump-sum transfers from the government, T_t; labor and capital income; proceeds from government bonds and profits from firms (the household owns all firms), Γ_t. The consumer uses these resources to consume, to pay consumption and income taxes, and to invest in government bonds. Its time t budget constraint is thus the following:

$$(1+\tau_t^c)c_t + \sum_{l=1}^{L} \frac{Q_t^l B_t^l}{P_t} + i_t = A_t, \tag{1.2}$$

$$A_t \equiv (1-\tau_t^n)\frac{W_t n_t}{P_t} + (1-\tau_t^k) r_{k,t} u_t K_{t-1} + \tau_t^k \delta_0 k_{t-1}^b + \sum_{l=1}^{L} \frac{Q_t^{l-1} B_{t-1}^l}{P_t} + T_t + \Gamma_t.$$

where A_t is the total wealth of the household in period t. Q_t^l, B_t^l are the nominal price and quantity of bonds with maturity l. W_t, $r_{k,t}$, P_t, τ_t^c, τ_t^n and τ_t^k are, respectively, the nominal wage, the rental rate of capital, the price index, the tax rates on consumption, on labor income and on capital income. Note that $Q_t^0 = 1$; so, bonds are sold at discount and pay one dollar upon maturity.

Notice that the capital income tax is levied on the rental rate times the capital service, $u_t K_{t-1}$, where u_t is the utilization rate of capital. Following Fernández-Villaverde et al. (2011), I allow the consumer to receive a depreciation allowance for the book value of capital, k_{t-1}^b. In addition, the representative consumer faces quadratic investment adjustment costs. The laws of motion of the capital stock and of the book value of capital are given by:

$$K_t = \left[1 - \delta_0 - \delta_1(u_t - 1) - \frac{\delta_2}{2}(u_t-1)^2\right] K_{t-1} + \left[1 - \frac{\kappa_i}{2}\left(\frac{i_t}{i_{t-1}} - 1\right)^2\right] i_t \tag{1.3}$$

$$k_t^b = (1-\delta_0) k_{t-1}^b + i_t \tag{1.4}$$

Note that the economic depreciation rate of capital, $\delta_0 + \delta_1(u_t-1) + \frac{\delta_2}{2}(u_t-1)^2$, is a quadratic function of capital utilization rate, meaning that the more capital is used beyond its normal intensity, the more it depreciates and leads to investment needs. I introduce depreciation allowances with the aim to reflect the actual U.S. tax system. This system uses the book value of capital instead of its replacement cost, and assumes that capital is used with normal intensity. This is why δ_0 is used as the depreciation rate in the accumulation process of the book value of capital (Equation (1.4)). The U.S. tax system allows to deduce the replacement of firms' depreciated capital from the tax base, therefore

[5] Standard preferences are obtained in this model by setting $\gamma = \frac{1}{\psi}$.

allowing them to reduce the distortions created by the tax system. Fernández-Villaverde et al. (2011) show that incorporating this feature into the model helps explain the size of the response of agents to fiscal uncertainty shocks.

A crucial feature of the model is investment adjustment costs (see the last term in Equation (1.3)). Apart from the more realistic and comprehensive definition of policy uncertainty, this is another crucial departure of my model from previous models that have dealt with the effect of policy risk on yields and premia. As will be shown in Section 1.4.2, responses of higher maturity yields are considerably influenced by whether the model accounts for investment adjustment costs. Bloom (2009) stresses that ignoring these adjustment costs leads to substantial bias in the effect of uncertainty shocks. I use the specification of investment adjustment costs proposed by Christiano et al. (2005), which is now widely accepted in the New-Keynesian literature. Eberly et al. (2012) show that this specification predicts the presence of a lagged-investment effect, a salient stylized fact on U.S. firms according to which the best predictor of current investment if lagged investment.

The consumer invests i_t in physical capital, and chooses c_t, n_t, $\{B_t^l\}_{l=1}^L$, K_t, u_t and k_t^b to maximize (1.1) subject to (1.2), (1.3) and (1.4). The first-order conditions of this problem can be written as:

$$\frac{\nu(1-\gamma)}{\theta} c_t^{\frac{\nu(1-\gamma)}{\theta}-1} (1-n_t)^{\frac{(1-\nu)(1-\gamma)}{\theta}} = \lambda_t(1+\tau_t^c), \tag{1.5}$$

$$\frac{1-\nu}{\nu} \frac{(1+\tau_t^c)c_t}{(1-\tau_t^n)(1-n_t)} = w_t, \tag{1.6}$$

$$Q_t^l = \beta E_t \left[\frac{\lambda_{t+1}}{\lambda_t} \left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}} \right)^{1-\frac{1}{\theta}} \frac{Q_{t+1}^{l-1}}{\pi_{t+1}} \right], \; l = 1, \cdots, L, \tag{1.7}$$

$$r_{k,t}(1-\tau_t^k) = q_t \left[\delta_1 + \delta_2(u_t - 1) \right], \tag{1.8}$$

$$\begin{aligned} q_t = \beta E_t \Bigg\{ & \frac{\lambda_{t+1}}{\lambda_t} \left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}} \right)^{1-\frac{1}{\theta}} \times \\ & \left[r_{k,t+1}(1-\tau_{t+1}^k)u_{t+1} + q_{t+1} \left(1 - \delta_0 - \delta_1(u_t - 1) - \frac{\delta_2}{2}(u_t-1)^2 \right) \right] \Bigg\}, \end{aligned} \tag{1.9}$$

$$q_t^b = \beta E_t \left\{ \frac{\lambda_{t+1}}{\lambda_t} \left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}} \right)^{1-\frac{1}{\theta}} \left[\delta_0 \tau_{t+1}^k + (1-\delta_0) q_{t+1}^b \right] \right\}, \tag{1.10}$$

$$q_t^b + q_t\left[1 - \frac{\kappa_i}{2}\left(\frac{i_t}{i_{t-1}} - 1\right)^2 - \kappa_i\frac{i_t}{i_{t-1}}\left(\frac{i_t}{i_{t-1}} - 1\right)\right] - 1$$

$$= \beta\kappa_i E_t\left\{\frac{\lambda_{t+1}}{\lambda_t}\left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}}\right)^{1-\frac{1}{\theta}}\left(\frac{i_{t+1}}{i_t}\right)^2\left(\frac{i_{t+1}}{i_t} - 1\right)q_{t+1}^b\right\}, \quad (1.11)$$

where λ_t is the Lagrange multiplier associated with the consumer's budget constraint, w_t is the real wage and $\pi_{t+1} = P_{t+1}/P_t$ is the gross inflation rate between t and $t+1$. (1.6) indicates the tradeoff between consumption and leisure, and (1.7) is the asset-pricing equation that is affected by recursive preferences through the term $\left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}}\right)^{1-\frac{1}{\theta}}$. q_t is the marginal Tobin's Q and q_t^b is the normalized multiplier of the book value of capital.

1.2.2 The Production Sector

There is a continuum of intermediate good firms (indexed by $i \in (0,1)$) that produce differentiated goods using labor and capital supplied by the representative household, and a final good firm that uses intermediate goods in order to produce a composite good. The final good producer evolves in a perfectly competitive environment and purchases $y_{i,t}$ units of each intermediate good to produce y_t units of the final good according to the following technology:

$$y_t = \left(\int_0^1 y_{i,t}^{\frac{\sigma-1}{\sigma}} di\right)^{\frac{\sigma}{\sigma-1}}. \quad (1.12)$$

The demand for the intermediate good $y_{i,t}$ is derived by the final good producer by minimizing its costs subject to (1.12) and taking prices as given. That minimization problem yields:

$$y_{i,t} = \left(\frac{P_{i,t}}{P_t}\right)^{-\sigma} y_t, \quad (1.13)$$

where $P_{i,t}$ is the price of the intermediate good i and σ is the elasticity of substitution between goods. The zero-profit condition implies that the price of the final good firm is:

$$P_t = \left(\int_0^1 P_{i,t}^{1-\sigma} di\right)^{\frac{1}{1-\sigma}}.$$

Each intermediate good firm i has monopoly power. It hires $n_{i,t}$ units of labor and buys $k_{i,t}$ units of capital, and sets its price, $P_{i,t}$ so as to maximize its profit. Its production technology is given by:

$$y_{i,t} = a_t k_{i,t}^{\alpha} n_{i,t}^{1-\alpha}, \quad (1.14)$$

where a_t is a technology shock that has the following process:

$$\log a_t = \rho_a \log a_{t-1} + \exp(\sigma_a)\varepsilon_{at},\ \ \varepsilon_{at} \sim N(0,1). \tag{1.15}$$

Many papers in the literature that specify stochastic volatility of technology shocks in order to achieve various goals (see Caldara et al. (2012), Born & Pfeifer (2014), Fernández-Villaverde et al. (2010), among others). Bansal & Yaron (2004) stress that including fluctuating economic uncertainty allows models such as ours to account for asset prices. However, I show that unlike recursive preferences, stochastic volatility of technology shocks are not crucial in order for our model to get asset prices right (see Section 1.4.1).
The intermediate good firm i also faces quadratic price adjustment costs à la Rotemberg (1982), given by:

$$AC_{i,t} = \frac{\kappa_p}{2}\left(\frac{P_{i,t}}{P_{i,t-1}} - \pi\right)^2 y_{i,t}, \tag{1.16}$$

π being the steady-state inflation rate. This adjustment cost, which increases with the size of the change vis-à-vis the steady-state inflation and the output, discourages firms who would like to reset their prices by increasing their costs if they do so. Nominal rigidities such as price adjustment costs are crucial for the propagation of uncertainty shocks; Fernández-Villaverde et al. (2011) show that they produce results similar to those obtained by Bloom (2009) after accounting for investment irreversibilities at the individual firm level.
The intermediate good firm maximizes:

$$E_t \sum_{s=0}^{\infty} M_{t+s}\left(\frac{P_{i,t+s}}{P_{t+s}} y_{i,t+s} - mc_t y_{i,t+s} - AC_{i,t+s}\right), \tag{1.17}$$

subject to (1.13) and (1.16), where mc_t is the real marginal cost of the firms obtained by minimizing its inputs costs; this minimization problem yields:

$$mc_t = mc_{i,t} = \frac{1}{a_t}\left(\frac{r_{k,t}}{\alpha}\right)^{\alpha}\left(\frac{w_t}{1-\alpha}\right)^{1-\alpha}. \tag{1.18}$$

The solution of the profit maximization problem is the following New-Keynesian Phillips curve taken at the symmetric equilibrium:

$$\left[(1-\sigma) + \sigma mc_t - \kappa\pi_t(\pi_t - \pi) + \frac{\sigma\kappa}{2}(\pi_t - \pi)^2\right] + \kappa E_t M_{t+1}\pi_{t+1}(\pi_{t+1} - \pi)\frac{y_{t+1}}{y_t} = 0. \tag{1.19}$$

In (1.17) and (1.19), M_t is the representative household's nominal stochastic discount factor defined by (1.20).

$$M_{t+1} = \beta\frac{\lambda_{t+1}}{\lambda_t}\left(\frac{V_{t+1}^{1-\gamma}}{E_t V_{t+1}^{1-\gamma}}\right)^{1-\frac{1}{\theta}}. \tag{1.20}$$

1.2.3 The Government

The government comprises a monetary authority and a fiscal authority. The monetary authority conducts monetary policy by setting the short-term nominal interest rate $R_t = (Q_t^1)^{-1}$ according to a standard Taylor rule that features policy uncertainty:

$$\log\left(\frac{R_t}{R}\right) = \rho_r \log\left(\frac{R_{t-1}}{R}\right) + (1-\rho_r)\left(\rho_{r,\pi}\log\left(\frac{\pi_t}{\pi}\right) + \rho_{r,y}\log\left(\frac{y_t}{y_{t-1}}\right)\right) + \exp(\sigma_{Rt})\varepsilon_{Rt}. \tag{1.21}$$

where $\varepsilon_{Rt} \sim N(0,1)$. In (1.21), the nominal interest rate reacts to its lag value (with an interest rate smoothing parameter ρ_r), and to inflation and output growth with the respective elasticities $\rho_{r,\pi}$ and $\rho_{r,y}$. ε_{Rt} is a monetary policy innovation. Variables without time subscripts are steady-state values. The standard deviation of the monetary policy shock is not constant as usually assumed in the literature. This is one of the main features of this model; by assuming time-variation for the standard deviation of the monetary policy innovation, the model accounts for monetary policy uncertainty.

The fiscal authority issues government bonds and levies taxes on consumption expenses, and on labor and capital income in order to finance exogenously determined government spending, g_t, depreciation allowances paid to the household, and lump-sum transfers to the representative consumer. The government budget constraint is given by:

$$\tau_t^c c_t + \sum_{l=1}^{L} \frac{Q_t^l B_t^l}{P_t} + \tau_t^n w_t n_t + \tau_t^k r_{k,t} u_t K_{t-1} = \sum_{l=1}^{L} \frac{Q_t^{l-1} B_{t-1}^l}{P_t} + T_t + g_t + \tau_t^k \delta_0 k_{t-1}^b, \tag{1.22}$$

where the government spending, g_t and the tax rates (τ_t^c, τ_t^n and τ_t^k) are the instruments that have laws of motion very close to those of Fernández-Villaverde et al. (2011). In their paper, each fiscal instrument reacts to its own lag, to lagged detrended output and to the lag of the deviation of the debt ratio vis-à-vis its targeted level. The fiscal rules that I specify below are slightly different from theirs since I do not include the last term. The motivation is that in the estimations of Fernández-Villaverde et al. (2011), almost all the coefficients associated with that term are not significantly different from zero; providing empirical evidence that fiscal instruments generally do not react to the deviation of the debt ratio vis-à-vis its targeted level. The fiscal rules are thus the following:

$$x_t - x = \rho_x(x_{t-1} - x) + \rho_{x,y}(y_{t-1} - y) + \exp(\sigma_{x,t})\varepsilon_{x,t},\ \varepsilon_{x,t} \sim N(0,1). \tag{1.23}$$

where $x \in \{g, \tau^c, \tau^n, \tau^k\}$. In this formula, each instrument reacts to its lag value and to deviation of the output from the steady state value. Here also, the fiscal rules are set to feature policy uncertainty. The processes of the standard deviations of the fiscal and monetary policy shocks, $\varepsilon_{z,t}$, are the same as in Fernández-Villaverde et al. (2011), Born

& Pfeifer (2014).

$$\sigma_{z,t} = (1-\rho_{\sigma_z})\sigma_z + \rho_{\sigma_z}\sigma_{z,t-1} + (1-\rho_{\sigma_z}^2)^{1/2}\eta_z \varepsilon_{\sigma_z,t},\ \varepsilon_{\sigma_z,t} \sim N(0,1) \ and \ \tau_z > 0. \quad (1.24)$$

where $z \in \{x, R\}$. σ_z is the unconditional mean of the corresponding standard deviation; it represents the average standard deviation of the policy innovation. η_z is the unconditional standard deviation of $\sigma_{z,t}$ and ρ_{σ_z} is its persistence. In this model, all shocks are independent vis-à-vis one another.

1.2.4 Aggregation and Equilibrium

In equilibrium, the budget constraint of the household and of the government, and the period-by-period profit yield the economy-wide resource constraint given by:

$$c_t + i_t + g_t + \frac{\kappa}{2}(\pi_t - \pi)^2 y_t = y_t, \quad (1.25)$$

where

$$y_t = a_t k_t^{\alpha} n_t^{1-\alpha},$$

$$n_t = \int_0^1 n_{i,t} di,$$

and

$$k_t = u_t K_{t-1} = \int_0^1 k_{i,t} di. \quad (1.26)$$

Definition: A *competitive equilibrium* is given by allocations $\mathcal{P}^i = \{n_{i,t}, k_{i,t}, P_{i,t}\}_{t=0}^{\infty}$ for each firm $i \in (0,1)$ and $\mathcal{C} = \{c_t, n_t, K_t, k_t^b, i_t, u_t, (B_t^l)_{l=1}^L\}_{t=0}^{\infty}$ for the representative household, such that, given a feasible government policy $\mathcal{G} = \{g_t, T_t, (B_t^l)_{l=1}^L\}_{t=0}^{\infty}$, a tax plan that satisfies (1.22) and a price system $\mathcal{Q} = \{\pi_t, w_t, r_{k,t}, (Q_t^l)_{l=1}^L\}_{t=0}^{\infty}$, the household maximizes (1.1) subject to (1.2), (1.3) and (1.4), and firms maximize (1.17) subject to (1.13) and (1.16).

1.3 Interest Rates, Term Premiums, Calibration and Solution Method

The main focus of this paper is on the dynamic effects of uncertainty risk on yields and on term premiums charged by traders. Following the literature, I define the gross yield on an $l-$period bond as

$$R_t^l = \left(Q_t^l\right)^{-\frac{1}{l}} \quad l = 1, \cdots, L.$$

There are many definitions of the term premium in the literature but the one I choose is the *holding term premium*, hp_t^l, defined as the difference between the expected holding period return on a $l-$period bond (i.e. the expected return on holding a $l-$period bond for only one period), and the return on a $1-$period bond[6]:

$$hp_t^l = E_t\left(\frac{Q_{t+1}^{l-1}}{Q_t^l}\right) - R_t^1 \tag{1.27}$$

The intuitive behind the definition of this term premium is very easy to understand. The holding term premium is the compensation to traders for holding a risky asset for one period instead of a risk-free bond (the one-period bond). Holding a $l-$period ($l > 1$) bond for only one period is risky because the price at which it could be redeemed in a future period is unknown.

Notice that the holding term premium defined in (1.27) is not ad-hoc since it can be derived from the asset-pricing equation (1.7) written as[7]

$$Q_t^l = Q_t^1 E_t\left(Q_{t+1}^{l-1}\right) + cov_t\left(Q_{t+1}^{l-1}, \frac{M_{t+1}}{\pi_{t+1}}\right).$$

Rearranging this equation yields the holding term premium

$$hp_t^l = E_t\left(\frac{Q_{t+1}^{l-1}}{Q_t^l}\right) - \frac{1}{Q_t^1} = -cov_t\left(\frac{Q_{t+1}^{l-1}}{Q_t^l}, \frac{M_{t+1}}{\pi_{t+1}}\frac{1}{Q_t^1}\right).$$

The term premium, thus, depends on the current spot rate and the conditional covariance between the holding period return and the real stochastic discount factor. In order to interpret the sign of the term premium, notice that:

$$M_{t+1} = \frac{\partial V_t/\partial A_{t+1}}{\partial V_t/\partial A_t}. \tag{1.28}$$

where, remember, A_t is the total wealth of the household in period t.

The above covariance is thus negative when the consumer is willing to increase its future wealth (i.e. when the marginal value of future wealth is high or the marginal value of current wealth is low and the agent would like to save for the future) but financial markets do not permit to do so, because the holding period return is too small. In this case, the consumer requires a positive term premium as a compensation. Furthermore, a positive term premium means that buying an l-period bond is cheaper than buying a one-period bond today and an $(l-1)$-period bond tomorrow[8]; for this reason, the yield curve is upward sloping. The term premium is negative when the holding period return covaries positively with the marginal value of future wealth, meaning that assets best pay off when the consumer's marginal utility of consumption is low; in this case the yield curve is downward sloping.

[6]Throughout the paper I use "*risk premium*" and "*term premium*" interchangeably although the latter is more accurate.

[7]Notice that $Q_t^1 = E_t\left(\frac{M_{t+1}}{\pi_{t+1}}\right)$

[8]To see this, note that $hp_t^l > 0$ implies that $Q_t^l - E_t\left(Q_{t+1}^{l-1}\right) < 0$.

The model is calibrated using parameters of the U.S. economy mostly taken from the literature; these parameters are estimates based quarterly data. So, one time period in the model represents a quarter. Table 1.1 summarizes the parameter values.

Table 1.1: Calibrated parameters

Parameter	Value	Description/Motivation/Source
Standard parameters		
α	0.33	Labor share of 67%
β	0.991	Steady state real interest rate of 3.68%
γ	35	Chosen to match the mean and upward slope of the U.S. yield curve
ψ	0.05	Chosen to match the mean and upward slope of the U.S. yield curve
δ_0	0.025	Steady-state depreciation rate
δ_1	0.0341	Ensures a steady-state capital utilization rate of 1
δ_2	0.0001	Fernández-Villaverde et al. (2011)
ν	0.3964	Ensures a steady-state labor supply of 1/3
κ_p	235.75	Fernández-Villaverde et al. (2011)
κ_i	3	Fernández-Villaverde et al. (2011)
σ	21	5% markup over the marginal cost at the steady state
π	1.005	Steady state annual inflation target of 2%
ρ_a	0.95	King & Rebelo (1999)
σ_a	-5.349	Born & Pfeifer (2014)
Monetary policy rule (Born & Pfeifer (2014))		
ρ_r	0.4889	Interest rate smoothing parameter
$\rho_{r,\pi}$	1.9691	Reaction of interest rate to inflation
$\rho_{r,y}$	1.2195	Reaction of interest rate to output growth
σ_R	-5.188	Unconditional mean of $\sigma_{R,t}$
ρ_{σ_r}	0.921	Persistence parameter of $\sigma_{R,t}$
η_R	0.934	Controls the standard deviation of of $\sigma_{R,t}$
Fiscal rules (Fernández-Villaverde et al. (2011))		
g	0.066	Unconditional mean of g_t
ρ_g	0.971	AR(1) coefficient of g_t
$\rho_{g,y}$	-0.009	Reaction of g_t to output
σ_g	-6.144	Unconditional mean of $\sigma_{g,t}$
ρ_{σ_g}	0.9251	AR(1) coefficient of $\sigma_{g,t}$
η_g	0.1804	Controls the standard deviation of $\sigma_{g,t}$
τ^c	0.0775	Unconditional mean of τ_t^c
ρ_{τ^c}	0.9946	AR(1) coefficient of τ_t^c
$\rho_{\tau^c,y}$	0.0023	Reaction of τ_t^c to output
σ_{τ^c}	-7.107	Unconditional mean of $\sigma_{\tau^c,t}$
$\rho_{\sigma_{\tau^c}}$	0.6248	AR(1) coefficient of $\sigma_{\tau^c,t}$
η_{τ^c}	0.6017	Controls the standard deviation of $\sigma_{\tau^c,t}$
τ^n	0.2244	Unconditional mean of τ_t^n
ρ_{τ^n}	0.9919	AR(1) coefficient of τ_t^n
$\rho_{\tau^n,y}$	0.0709	Reaction of τ_t^n to output
σ_{τ^n}	-6.005	Unconditional mean of $\sigma_{\tau^n,t}$
$\rho_{\sigma_{\tau^n}}$	0.301	AR(1) coefficient of $\sigma_{\tau^n,t}$
η_{τ^n}	0.9454	Controls the standard deviation of $\sigma_{\tau^n,t}$
τ^k	0.3712	Unconditional mean of τ_t^k
ρ_{τ^k}	0.9668	AR(1) coefficient of τ_t^k
$\rho_{\tau^k,y}$	0.1005	Reaction of τ_t^k to output
σ_{τ^k}	-4.962	Unconditional mean of $\sigma_{\tau^k,t}$
$\rho_{\sigma_{\tau^k}}$	0.7659	AR(1) coefficient of $\sigma_{\tau^k,t}$
η_{τ^k}	0.5758	Controls the standard deviation of $\sigma_{\tau^k,t}$

The parameters of the monetary policy rule are estimated by Born & Pfeifer (2014)[9], and those of the fiscal rules are estimated by Fernández-Villaverde et al. (2011). The remaining parameters have values that are standard and widely used in the literature. The capital share α is set to 0.33, β is set to 0.991 to ensure a steady-state real short-term interest rate of 3.68%. ψ is set to 0.05; a value that is popular in this literature and falls in the range of values estimated by Ruge-Murcia (2017). The risk aversion parameter γ is set to 35, a value closed to the estimates of Ruge-Murcia (2017). Together, the values of ψ and γ allow me to match the upward slope and the mean of the U.S. yield curve for the period 1970Q1-2010Q2[10]. Most of the papers in the literature require a large value of γ in order for the model to replicate asset-prices features (see for instance Rudebusch & Swanson (2012), Caldara et al. (2012), among others). This large value, however, amplifies the precautionary behavior of agents; it is nevertheless plausible since it belongs to the range of estimates of the risk aversion parameter in the literature; moreover, the results presented below are robust to the choice of γ (see for instance panel A of Figure 1.7). ν is set to 0.3964 to ensure a steady-state labor supply of 1/3. The price and investment adjustment cost parameters, κ_p and κ_i, are respectively set to 235.75 and 3 as in Fernández-Villaverde et al. (2011). σ is set to 21 so that intermediate good firms have a 5% markup over their marginal cost at the steady state. The steady-state level of inflation is set to $1.02^{0.25}$ to have most central banks' 2% inflation target at the steady-state. δ_0 and δ_2 are taken from Fernández-Villaverde et al. (2011) and δ_1 is set to 0.0341 to ensure a steady-state utilization rate of capital of 1. Finally, the persistence parameter of the technology shock, ρ_a, is set to 0.95 and its standard deviation, σ_a, is set to -5.349; these are standard values in the literature.

The model is solved using a third-order perturbation around the deterministic steady state (an extension of the second-order perturbation method by Schmitt-Grohé & Uribe (2004)) since there is not a closed-form solution due to the highly nonlinear nature of the model, and since there is a high number of state variables making it extremely difficult to use global solution methods. A third-order approximation to the policy function is particularly important because, as stressed by Fernández-Villaverde et al. (2010), innovations to volatility shocks appear in a time-varying manner only in the third-order terms of the Taylor approximation of the policy function. Fernández-Villaverde et al. (2010), Ruge-Murcia (2012), Andreasen et al. (2013) show how to use third-order approximation in order to estimate DSGE models. The latter provides a pruning algorithm that helps to avoid the explosive sample paths that higher-order approximations usually generate; I apply this algorithm in my simulations.

[9] Born & Pfeifer (2014) define the process of the standard deviation of the monetary policy as $\sigma_{R,t} = (1-\rho_{\sigma_R})\sigma_R + \rho_{\sigma_R}\sigma_{R,t-1} + \eta_R^{'}\varepsilon_{\sigma_R,t}$. So, $\eta_R = \frac{\eta_R^{'}}{(1-\rho_{\sigma_R}^2)^{1/2}}$.

[10] Although these values of ψ and γ may seem very different from those used in the standard DSGE literature, the values used here are similar to the ones estimated in many papers in the asset pricing literature. See Ruge-Murcia (2017) and references therein for a number of other papers with values closed to the ones used here.

1.4 Theoretical Model Results

This section presents the results of my analyses. First of all, I present the contribution of policy uncertainty to bond premiums and decompose it to investigate the effects of each source of the uncertainty (monetary and fiscal) on bond yields and premiums; these results are based on simulated observations computed after solving the model using third-order perturbation around the deterministic steady-state. Then, I perform impulse response analyses in order to document the dynamic effects of policy risk shocks on bond yields. A section is also dedicated to showing the role played by investment adjustment costs. Some sensitivity analyses are performed in order to have a better understanding of the way policy risk transmits to yields, and to show the parameters of the model that most account for the observed dynamics. In spite of the fact that I choose some parameter value to match important characteristics of the yield curve, the numbers presented in this section are only indicative since almost all of the model's parameters are calibrated; the qualitative results, however remain robust (i.e. they do not depend on calibration).

1.4.1 Contribution of Policy Uncertainty to Bonds Yields and premiums

Getting the Yield Curve Right

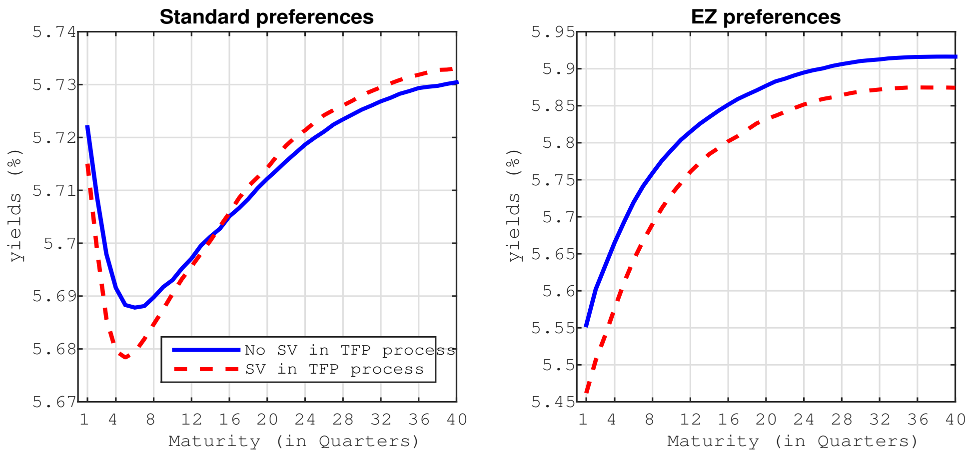

Figure 1.1: Importance of recursive preferences in getting the yield curve right (model without policy uncertainty).

Notes: The solid blue line in the left (resp. right) panel represents the yield curve in the model featuring standard (resp. recursive) preferences and homoskedastic technology; and the dashed red line in the same panel depicts the yield curve in the model featuring standard (resp. recursive) preferences and stochastic volatility to the technology shock.

Many papers in the term structure literature have found standard DSGE models unable to fit the macro side of the model and, at the same time, replicate some major features of asset prices (see for instance Backus et al. (1989) and den Haan (1995) among many others). Bansal & Yaron (2004) stressed that recursive preferences combined with stochastic volatility and long-run risks should help DSGE models account for asset prices. Rudebusch & Swanson (2012) also combine recursive preferences and long-run risks to get to a similar conclusion. One specific problem in this literature is the difficulties that standard DSGE models have to replicate the upward slope of the yield curve. In our model, the presence of stochastic volatilities does not seem to matter much for fixing this problem, unlike recursive preferences. Indeed, as shown in the left panel of Figure 1.1, the model fails to replicate the upward slope of the yield curve when standard preferences are used; particularly the short-end of the yield curve slopes downward. The right panel of Figure 1.1 shows that the yield curve slopes upward no matter if the model features stochastic volatility (solid blue line) or not (dashed red line). Note that to be consistent with the main objective of this paper, we need to get the yield curve right in a version of the model that does not feature policy uncertainty. For this reason, the yield curves presented in Figure 1.1 are computed using a version of the model where the standard deviations of policy instruments are replaced with their unconditional means[11].

In order to have a better understanding of the role played by recursive preferences, I simulate data from two versions of the model, one featuring recursive preferences and not the other, and I compare the correlations between current consumption and expected holding period returns in these two models. The intuition is that if the yield curve was upward sloping, this correlation would be negative (see discussion in Section 3.3). The idea is therefore to assess the role of recursive preferences in getting the sign of the correlation right. Figure A.1 in Appendix B.2 shows that recursive preferences do play a very important role; the graph shows the correlation between c_t and $E_t\left(\frac{Q_{t+1}^{l-1}}{Q_t^l}\right)$ as a function of maturity l. The orange dashed line (right axis) shows that this correlation is always negative when households' preferences are recursive. In contrast, the solid blue line of the figure (left axis) shows that the model fails to generate this negative correlation when I use standard preferences.

Policy Uncertainty and Yields

The main question that I aim to answer in this section is:*how do yields behave when the economy is constantly hit by unpredictable shocks to the volatility of policy instruments?* Figure 1.2 shows the effect of policy uncertainty on yields, and the contribution of our two main sources of policy uncertainty (monetary and fiscal) to this effect[12]. The main observation is that policy uncertainty decreases yields. The top left panel shows the yield curve generated by a version of the model that does not include policy uncertainty (solid black line), and the one generated by a model that includes both monetary and fiscal policy uncertainties; it emerges that policy uncertainty shifts the level of the curve

[11]The blue lines in Figure 1.1 show the median yield curve ifor the model the standard deviation of the technology shock is stochastic whereas it is not in the bottom panels. Meaning that in the upper panels I replace σ_a in Equation (1.15) by $\sigma_{a,t}$, and specify the stochastic volatility process of technology shocks as $\sigma_{a,t} = (1-\rho_{\sigma_a})\sigma_a + \rho_{\sigma_a}\sigma_{a,t-1} + \eta_a\varepsilon_{\sigma_a,t},\ \varepsilon_{\sigma_a,t} \sim N(0,1)$ *and* $\eta_a > 0$.

[12]In this figure, the solid black line depicts the yield curve in the model not featuring policy uncertainty.

downward and steepens its slope[13]. The level and the slope are respectively 5.75% and 0.50% when the model accounts for policy uncertainty compared to 5.84% and 0.36% when it does not. Economic policy uncertainty, thus, decreases yields' level by 8.56 basis points and increases the term spread by 13.54 basis points.

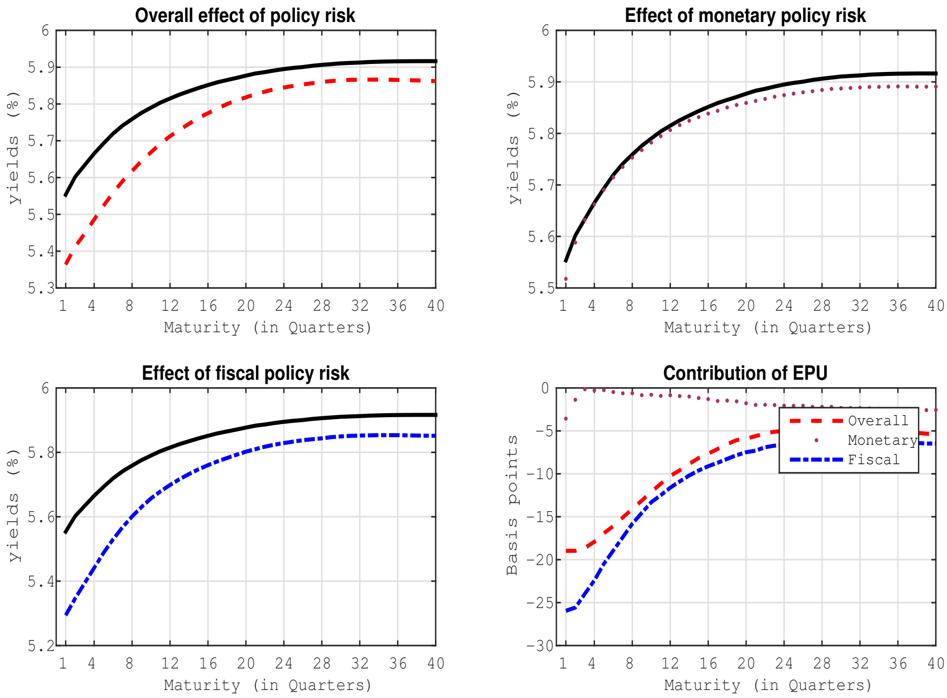

Figure 1.2: Effect/contribution of policy uncertainty to yields.

Notes: The solid black line depicts the yield curve in the model without policy uncertainty whereas the dashed red (resp. dotted brown, and dash-dotted blue) line depicts the effect/contribution of the joint (resp. monetary and fiscal) policy uncertainty.

In order to gain more insight into how policy uncertainty affects yields and compare my results to previous findings in the literature, I use two other versions of the model. The first one includes only monetary policy uncertainty and the second, only fiscal policy uncertainty; in each case, the standard deviation of the shock to the policy instrument not being considered for policy uncertainty is set to its unconditional mean. The top right panel of Figure 1.2 shows that monetary policy uncertainty decreases interest rates in general, but the decrease is more pronounced on the short-end of the yield curve. The level of the yield curve decreases by 1.68 basis points and its slope increases by 1.02 basis points. Fiscal policy uncertainty has a similar effect. It decreases the level of the yield curve by 10.55 basis points, and increases its slope by 19.47 basis points (bottom left panel).

The bottom right panel of Figure 1.2 shows the contribution of each source of policy

[13] The level of the yield curve is defined as the equally weighted average of bonds yields of maturity 1 to 40, and the slope is defined as the spread between the highest maturity bond yield and the lowest maturity bond yield, i.e $R^{40} - R^1$.

uncertainty to the yields; the contribution is defined as the difference in yields between the relevant version of the model and the model without policy uncertainty. It turns out that both sources decrease yields at all maturities but the effect of monetary policy uncertainty (purple dotted line) is smaller than that of fiscal policy uncertainty (blue dash-dotted line). Fiscal policy uncertainty affects the short end of the curve more than its long end, thus behaving like a slope factor. Monetary policy uncertainty, however, tends to have only very small effects on yields. A striking observation stemming from this analysis is that although both policy risks affect yields negatively, their effects do not add up to the overall effect of economic policy uncertainty on the yield curve; the overall effect even turns out to be smaller than the fiscal effect. Moreover this overall effect seems to be a non-linear combination of the monetary and the fiscal effects.

Policy Uncertainty and Term premiums

The model generates positive and hump-shaped term premiums and the highest term premium is reached at the 4-year maturity. Positive term premiums mean that for the representative consumer, the expected return on buying an $l(> 2)$-period bond and selling it in the next period is greater than that of a 1-period bond. This explains the upward slope of the yield curve presented above. Figure 1.3 shows the term premiums (in basis points) generated by the different versions of the model presented earlier. Here too, the black line depicts the graph of term premiums in the model that does not feature policy uncertainty; the red dashed line depicts the combined effect of monetary and fiscal policy uncertainty; the purple dotted line is that of monetary policy uncertainty alone and the blue dash-dotted line is that of fiscal policy uncertainty alone. The top left panel shows that term premiums are lower when the model features both monetary and fiscal policy uncertainty. One can observe that the negative effect of policy uncertainty on premiums is small for medium-term bonds (less than 0.1 basis points for the 4-year maturity bond as shown by the red dashed line in the bottom right panel of the figure), but this negative effect is higher for short and the long maturity bonds (Policy uncertainty increases term premiums on 3-month and ten-year bonds by almost 0.4 basis points). On average, economic policy uncertainty decreases term premia by 0.21 basis points. As the bottom left panel of Figure 1.3 shows, the effect of fiscal policy uncertainty is qualitatively similar to the overall effect but 0.03 basis points more amplified on average. In total contrast with the overall effect and the fiscal effect, monetary policy uncertainty increases term premiums on bonds and the effect is almost uniform across maturities (see the top left panel of the figure and the purple dotted line in the bottom right panel that shows an average positive effect 0.22 basis points). As for the case of yields, the overall effect seems to be a non-linear combination of the fiscal and the monetary effects.

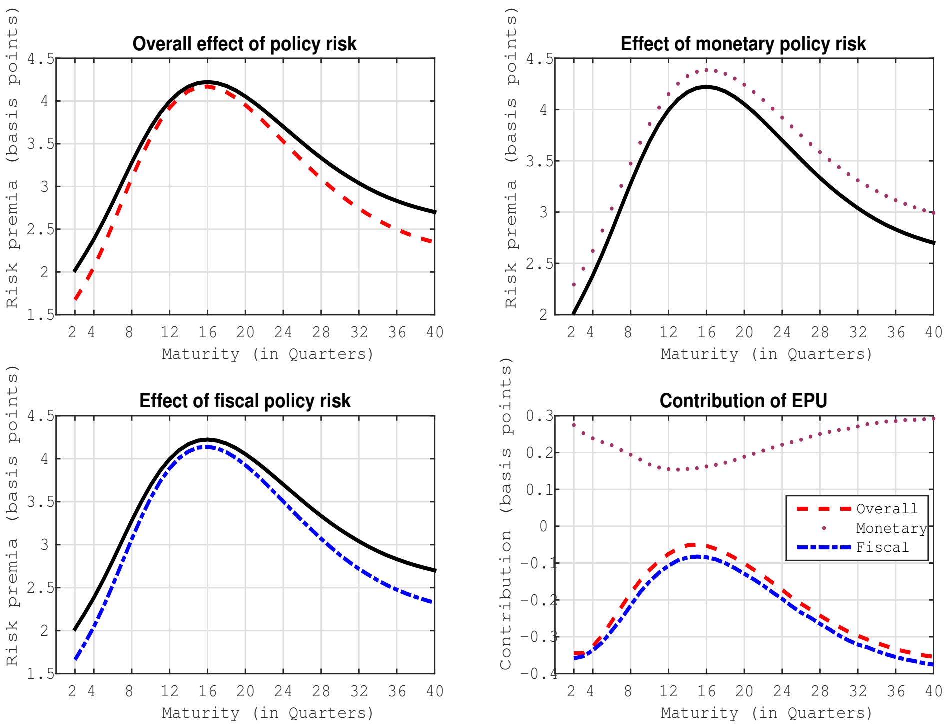

Figure 1.3: Effects of policy uncertainty on premiums (in basis points).

Notes: Black solid line: no policy uncertainty; purple dotted line: monetary policy uncertainty; blue dash-dotted line: fiscal policy uncertainty; red dashed line: overall.

Why do Yields and premiums Decrease?

In the results presented above, the economy is subject to both positive and negative policy risk shocks that have the same probability of occurrence. Intuitively, if negative and positive risk shocks had symmetric effects on yields, the overall effect of such shocks in our simulations would be nil on average. But given the degree of nonlinearity in the model, positive and negative shocks are very unlikely to have symmetric effects on yields and premiums. This motivates an additional analysis that compares the effect of positive and negative policy risk shocks on yields and premiums, on impact. Figure A.2 in Appendix B.2 shows that a positive two-standard deviation policy risk shock increases yields at all maturities and a negative shock of the same magnitude decreases yields. The effect of the negative shock is higher in absolute value[14]. Therefore, agents are more sensitive to a policy that tends to narrow the range of fluctuation of innovations to tax rates, government spending, and the monetary policy rate. The magnitude of yields decreases in such circumstances is higher than the magnitude of their increases when agents expect a spread in the fluctuation ranges of innovations to policy instruments; consequently yields decrease on average. An analogous explanation holds for term premiums.

I have shown that the effect of policy risk on yields and premiums is due to asymmetric

[14]Note that in Figure A.2, the red bars representing the effects of negative shocks are inverted in order to show the difference in the magnitude of both positive and negative shocks

impacts of positive and negative shocks, but the mechanism through which these shocks transmit to yields is presented in Section 1.4.2. To have a foretaste of the transmission mechanism, note that in periods of high policy uncertainty yields increase to compensate agents for real losses from bond proceeds due to higher expected inflation (as well as low consumption smoothing gains from bonds). In contrast, in periods of low policy uncertainty—which I have referred to as periods of higher transparency—inflation expectations are very low and the certainty equivalent future consumption is high, leading to decrease in yields that are more pronounced. As a consequence, yields fall by more (in absolute terms) than they increase under a high policy uncertainty environment. This leads to a lower average yield curve and lower term premiums.

1.4.2 Dynamic Effects

Impulse Response Analysis

I now turn to studying the dynamic effects of policy uncertainty on yields; I show the effects of shocks to the innovations of standard deviations of policy instruments. Figure 1.4 plots the impulse responses of various bond yields to a two-standard deviation policy risk shock. The responses represent deviations from the ergodic steady-state and are expressed in annualized basis points. Born & Pfeifer (2014) estimate historical volatilities of policy instruments and show that they co-move most of the time, particularly in periods of high economic uncertainty like during the Great Recession. Moreover, Fernández-Villaverde et al. (2011) stress that fiscal policy uncertainty can be parsimoniously captured by a simultaneous increase in the volatilities of the innovations to all fiscal instruments. Following these two papers and also what is common practice in the uncertainty literature, I define a policy risk shock as a simultaneous two-standard deviation increase in the volatility of innovations to R_t, g_t,τ_t^c, τ_t^n and τ_t^k, meaning that agents expect a simultaneous mean-preserving spread of the distribution of future monetary policy rates, government spending and tax rates. The black solid lines in Figure 1.4 depict the overall effects of the policy risk shock; the purple dotted lines represent the effects of a two-standard deviation monetary policy risk shock and the blue dashed lines represent the effects of a two-standard deviation fiscal policy risk shock. The effects of policy risk shocks on business cycle fluctuations are shown in Figure A.3, but they are not the subject of our main focus here. Note however that these effects are similar to the results of Born & Pfeifer (2014) and Fernández-Villaverde et al. (2011), which is not surprising given the similarities of our model economy to theirs.

Following a joint policy uncertainty shock, the 3-month interest rate increases by 20 basis points on impact, peaks at 29 basis points in the second period, and gradually decreases to become negative four years after the shock; this hump-shaped response appears only for the very short-end of the yield curve but the impact effects and the peaks are higher for the 6- and 9-month yields (not shown). More generally, the initial impact of the policy risk shock remains above 20 basis points for all yields of maturity 2-year and below, before

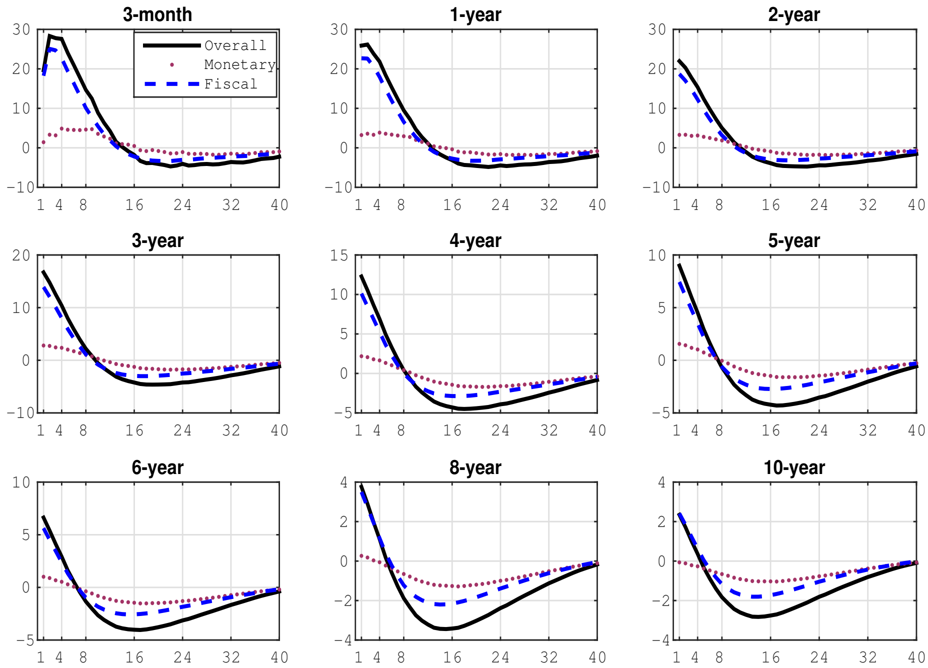

Figure 1.4: Impulse responses to a two-standard deviation policy risk shock (in annualized basis points).

Notes: The solid black lines depict the effects of a joint policy risk shock; the purple dotted lines depict the effects of the monetary policy risk shock, and the blue dashed line represent the effect of the fiscal policy risk shock. Horizontal axes represent quarters

gradually decreasing with maturity. The effect of the shock on medium and long rates does not feature a hump-shaped pattern as for the short rates; after the initial positive effects—that decrease with maturity (for instance 22, 5.8 and 2.2 basis points for 2-, 5- and 10-year bond yields respectively)—yields decrease and the effects also become negative before returning to the steady-state after roughly 40 quarters. Notice that the higher the maturity, the higher the rate at which yields decrease after the initial effect and, therefore, the smaller the number of periods after which the effects of the shock become negative. For instance, the effect of the shock on the 2-year rate turns negative after three years whereas it does so after only 1.5 years for the 10-year rate.

The effects of the joint policy risk shock is the combination of two effects. The monetary policy risk effect depicted by the purple dotted lines and the fiscal policy risk effect depicted by the blue dashed lines. The effects of these two sources of policy uncertainty appear to be qualitatively similar—at least for the short-end of the yield curve—and seem to contribute to the overall effect in a quasi-linear way[15]. However, this overall effect is mostly due to the fiscal effect as both effects move very closely to each other, the latter being just slightly lower than the former. Furthermore, the monetary effect appears to

[15]We did the sum of the fiscal and the monetary effects and the results are surprisingly very close to the overall effect.

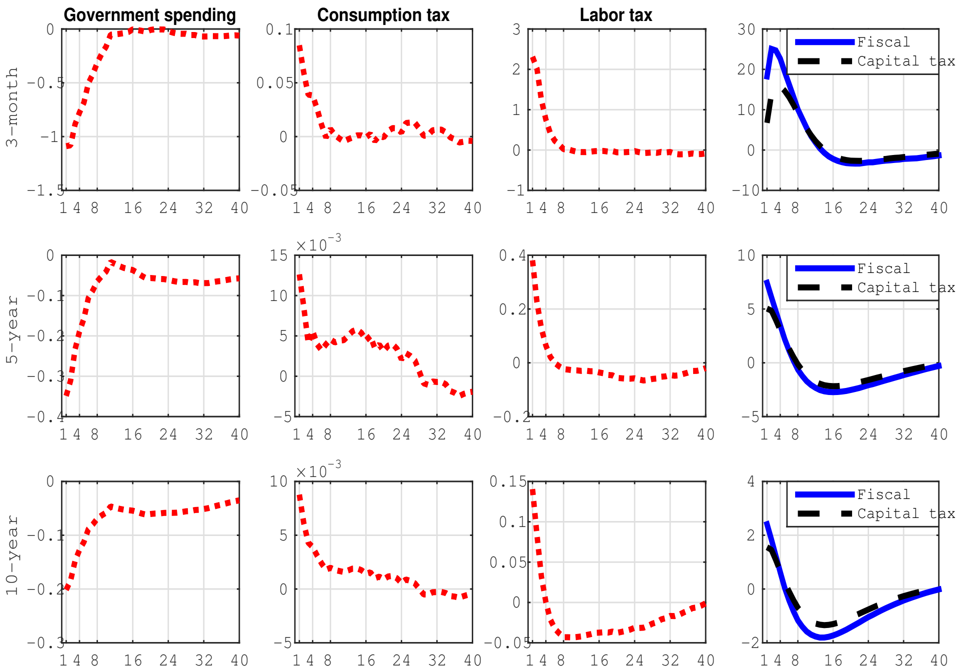

Figure 1.5: Decomposing the effects of fiscal policy risk by instrument.

Notes: The first three columns depict the effect of a two-standard deviation shock to the volatility of, respectively, government spending, consumption tax rate and labour tax rate. The solid blue line in the right panel (fourth row) represents the overall effect of the two-standard deviation fiscal policy risk shock and the black dashed lines represent the effect of a two-standard deviation shock to the volatility of capital tax rates.

be very small compared to the fiscal effect. Following a monetary policy risk shock, very short-term yields (one-year or less) increase in a hump-shaped manner, peaking at 5 and 4 basis points after one year respectively for the 3-month and the 1-year bond rates[16]. After the initial impact (which decreases with maturity and is nil for the 10-year rate), the effect decreases and becomes negative with roughly the same timing as for the overall effect. Variance decomposition shows that the monetary policy risk shock explains roughly 5% of the overall effect of the policy risk for the 3-month maturity bond, and this share gradually increases with maturity to reach almost 30% for the 10-year maturity bond.

In order to gain more insight about which sources of fiscal policy uncertainty account the most for the observed dynamic effects, I plot in Figure 1.5 the responses of yields to a policy risk shock related to each fiscal instrument. The first three columns of the figure depict the effect of a two-standard deviation shock to the volatility of, respectively, government spending innovations ($\varepsilon_{\sigma_g,t}$), consumption tax rate innovations ($\varepsilon_{\sigma_{\tau^c},t}$) and labour tax rate innovations ($\varepsilon_{\sigma_{\tau^n},t}$). The solid blue lines in the right panels (fourth column) represent the overall effects of the two-standard deviation fiscal policy risk shock

[16]Monetary policy uncertainty is therefore partly responsible for the hump-shaped response of the short-end of the yield curve.

(i.e. the blue dashed line in Figure 1.4) and the black dashed lines represent the effect of a two-standard deviation shock to the volatility of innovations to capital tax rates ($\varepsilon_{\sigma_{\tau^k},t}$). As it is very apparent, the shock to the volatility of innovations to capital tax rates accounts for almost all the fiscal effect and its contribution increases with maturity. As stressed by Fernández-Villaverde et al. (2011), this is because only capital taxes distort the intertemporal resource allocation in the model. A non-negligible part of the effects in the immediate aftermath of the shock is attributable to the government spending uncertainty shock. Unlike all other shocks that have an initial positive effect on yields, this shock has a negative effect albeit small compared to the "capital" effect (-1.2 basis points for the 3-month yields and -0.2 basis points for 10-year yields); this effect reduces the effect on impact of the joint shock and partly explains the hump-shaped responses of short rates presented earlier. An important observation that emerges from this investigation is that the contribution of the shock associated to the volatility of innovations to labour tax rates is very small but not negligible. In contrast, the contribution of the shock associated to the volatility of innovations to consumption tax rates is negligible.

Transmission Mechanism and the Role of Investment Adjustment Costs

In order to understand the mechanism underlying the effects of policy risk on yields, let's return to the asset-pricing equation (1.7) and investigate the effect that policy uncertainty has on relevant forward looking variables, namely expected inflation, and the marginal value of future wealth. According to equation (1.28), the determinants of households' asset-pricing decisions in each period are summarized in these two variables[17]. Simulations show that policy uncertainty affects yields through two channels: on the one hand, higher uncertainty leads to higher expected inflation (the "inflation channel"), due to higher future prices as firms expect future marginal costs to be more volatile in periods following the shock, and increase prices in order to avoid getting stuck with low prices given that they face price adjustment costs. The increase in prices occurs despite the reduction in real marginal costs (see Figure A.3), meaning that firms increase their markup in the aftermath of the shock. On the other hand, increased policy risk increases the volatility of consumption; to hedge against that, households increase their demand for bonds, putting downward pressures on yields (I refer to this as the "smoothing" channel). The inflation channel dominates; so, the fall in yields triggered by higher bond demand from households is outweighed by the increase in yields due to higher inflation risk premiums.

The smoothing channel is explained by the need for resource reallocation following the joint policy risk shock. Given that most of the distortions associated to the shock come from fiscal policy uncertainty (see Figure 1.4), and specifically from the capital tax policy uncertainty (see the right panels of Figure 1.5), the household would like to avoid these distortions by using means other than investment to transfer wealth intertemporally.

[17] An important notice here is that the marginal value of wealth equals the marginal value of consumption, up to a small value related to the consumption tax rate. Interpreting the results using either of the two is, therefore, not problematic.

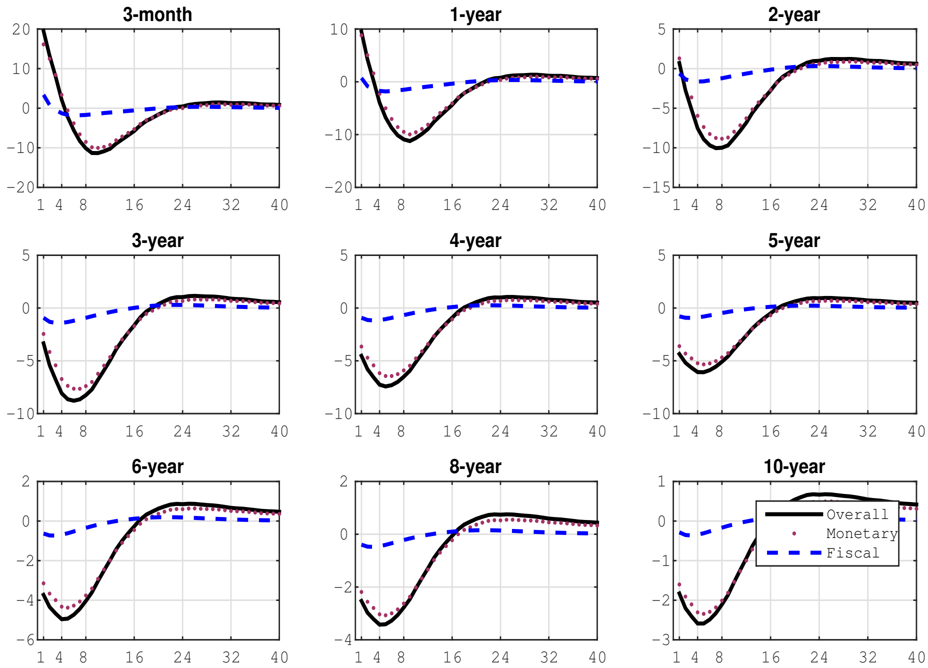

Figure 1.6: Impulse responses to a two-standard deviation policy risk shock (in annualized basis points) in the model without investment adjustment costs.

Notes: The solid black lines depict the effects of the overall policy risk shock; the purple dotted lines depict the effects of the monetary policy risk shock, and the blue dashed line represent the effect of the fiscal policy risk shock. Horizontal axes represent quarters

Bonds offer this possibility. However, investment adjustment costs prevent this resource reallocation (reducing investment and increasing demand for bonds) from being fully operative. This role of investment adjustment costs plays more on bonds of medium and high maturities.

Figure 1.6 illustrates very well this crucial role of investment adjustment costs for resource reallocation, but also for determining the importance of each source of uncertainty (monetary versus fiscal) in the response of medium and long yields. Reducing agents' investment adjustment costs improves their ability to transfer wealth intertemporally, enhancing the smoothing channel and reducing the magnitude of yields increase following a shock. Absent these costs[18], the smoothing channel becomes dominant in the medium- and long-run and the corresponding yields decrease. The intuition is that agents gradually reduce investment in response to the shock, but the reduction is more pronounced when they do not face adjustment costs. Indeed, the negative effect of the shock on the stock of capital ten years after, is more than three times higher without adjustment costs (See Figures A.3 and A.4 in Appendix B.2). When reducing investment, agents increase the utilization rate of capital in order to preserve capital revenues, but at the same time,

[18]This is achieved by setting $\kappa_i = 0$

increase their demand for medium- and long-term bonds; ultimately, the smoothing channel outweighs the inflation channel and yields on bonds of these maturities decrease (see rows 2 and 3 of Figure 1.6). A similar argument holds for all sources of policy risk.

Also noteworthy is the fact that investment adjustment costs shifts the relative importance of monetary and fiscal policy risk shocks for the response of yields on medium and long bonds. Absent these costs, the fiscal effect is small compared to the monetary effect. This suggests that investment adjustment costs are responsible for (or at least play an important role in) the subdued response of yields to monetary policy risk shocks. As emphasized in the previous section, the fiscal effect is channeled through uncertainty related to capital tax policy. This effect is amplified by adjustment costs to investment. Under flexible investment adjustment, adjusting to the distortions caused by uncertainty is no longer costly; which considerably dampens the fiscal effect. The opposite argument (regarding the role of investment adjustment costs) holds for the effect of monetary policy uncertainty on medium- and long-term bonds.

A positive monetary policy risk shock makes future interest rates and, hence, consumption more volatile (smoothing channel) while, at the same time, increasing inflation expectations (inflation channel). Without investment adjustment costs, the smoothing channel becomes the most important channel in the medium- and long-run. The reason is that resources originally dedicated to investment can costlessly be used to increase demand for bonds and hedge against future consumption volatility, while eventually managing capital utilization rates to meet capital demand from firms. This monetary effect is thus amplified by the absence of investment adjustment costs. The still dominant inflation effect on short-term yields is explained by very high inflation expectations, again due to the reaction of firms to the shock.

As just shown, investment adjustment cost play an important role for the sign of the response of medium and long yields to policy risk shocks. This finding is similar to that of Bloom (2009) on the importance of investment adjustment costs in accounting for the effects of uncertainty shocks. This main feature, together with the explicit modelling of the government behavior, lead my results on the effect of policy risk shocks on medium and long yields to be different from that of Leippold & Matthys (2015) who find a negative effect of policy uncertainty on yields. On the one hand and as emphasized in the Introduction, these authors include an exogenous process that supposedly captures the dynamics of government policy uncertainty; their process is a simple one that affects real output growth and productivity but is unrelated to any government action. There is no guaranty that such a process actually captures government policy uncertainty, and in fact given the way it enters the model, it cannot be distinguish from, say, innovation or human capital. The introduction of an actual government in our model leads to a significant difference in the way yields respond to government policy uncertainty. It switches the sign of the response of short yields which become positive and, as shown previously, this is due to the very high inflation risk premiums charged by investors on short-term bonds following a policy risk shock. On the other hand, the absence of investment adjustment costs in the model of Leippold & Matthys (2015) leads the effect the smoothing channel to be dominant, and that of the inflation channel to be subdued, explaining the difference in our results. The results presented in Section 1.5 show that the theoretical results presented above are in line with empirical fact, confirming that the inflation channel does play a very crucial role in the way yields respond to policy uncertainty shocks.

Sensitivity analyses

The results presented so far and the transmission mechanism of policy uncertainty surely depend on calibrated parameters of the model. This section presents how robust these results are to variations in some of these parameters; this is also useful for a deeper understanding of the transmission mechanism, and also to know which parameters play a significant role. The results of this exercise are presented in the panels of Figure 1.7 where I compare the baseline response of the 5-year bond rate to a standard policy risk shock (black solid line) to the responses when some parameters are changed. The title of each panel specifies the parameter whose value is modified in that particular panel; for instance, the title of panel A, "$A:\ \gamma = \{baseline;\ 2;\ 20\}$", means that the graphs plot the responses of the 5-year bond rate for three different values of parameter γ while keeping all other parameters unchanged: the baseline value (solid black line), $\gamma = 2$ (blue dashed line) and $\gamma = 20$ (red dash-dotted line).

In the first two panels of row one (A, E), I perturb the parameters driving household preferences. First, changing the risk aversion parameter does affect the response of yields (see panel A) meaning that this parameter plays a very important role in getting asset prices right (see Section 1.4.1) but does not influence agents' pricing of policy risk. This may be because what matters most for investors when the shock hits is their ability to so use bonds in order to smooth out consumption. Indeed, increasing the IES from the baseline value of 0.05 to 0.25 and 0.5 (panel B) amplifies the consumption smoothing channel and, as a consequence, the effect of the policy risk shock on yields is less reduced (the impact effect decreases from about 10 basis points to about 6 and 4 basis points respectively).

Panels D, E and F show how parameters controlling inflation volatility and firms pricing decisions affect the impact of policy risk on interest rates. First, note that a weak inflation response of the central bank ($\rho_{r,\pi} = 1.31$ in panel D) leads to higher inflation expectations and amplifies the inflation channel described above; consequently yields increase further. However, a strong inflation response ($\rho_{r,\pi} = 4$) reduces inflation expectations, dampens the inflation channel which is now compensated by the smoothing channel; in this case policy risk has almost no effect on yields. The behavior of the central bank towards inflation plays an important role in explaining the influence of the price markup parameter (σ) and the price adjustment cost parameter (κ_p). When firms set prices using a high markup over their marginal cost ($\sigma = 11$ in panel E), inflation increases and moves away from the central bank's target, triggering the dampening of the inflation channel discussed previously; this leads to a lower impact of the shock. In contrast, if firms set prices

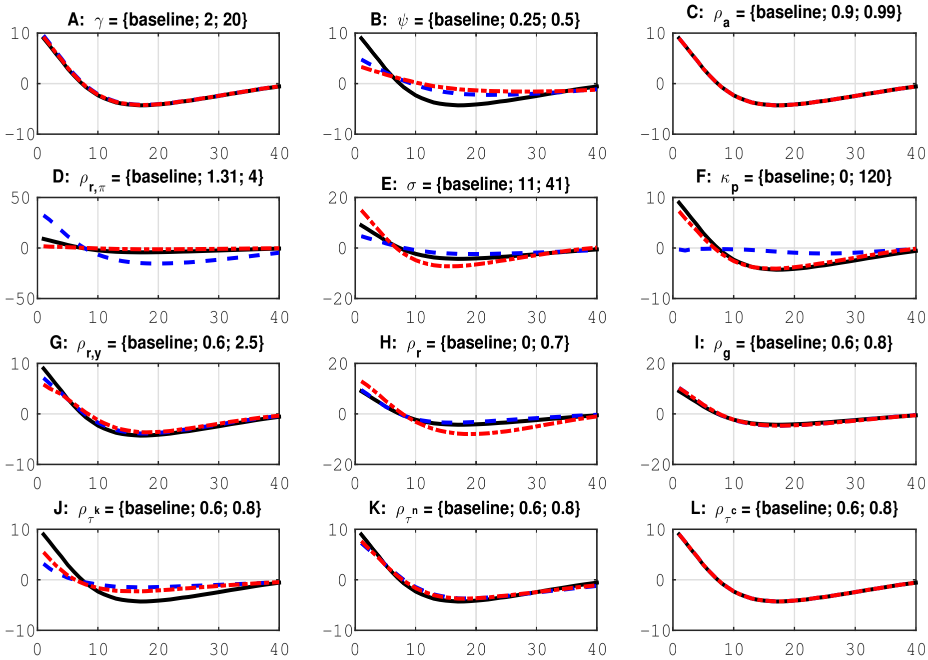

Figure 1.7: Response of the 5-year bond rate to a typical policy risk shock (in annualized basis points).

Notes: "X : $x = \{baseline;\ a;\ b\}$" in the title of panel X means that the graph plots the responses for three different values of parameter x: baseline (solid black line), a (blue dashed line) and b (red dash-dotted line).

using a low markup ($\sigma = 41$) the central may not have to intervene, leading agents to increase their inflation expectations and ask higher compensations to keep bonds in their portfolio. Panel F shows that price stickiness matters but not its degree; the response of the 5-year bond rate to the policy risk shock does not change much when the price adjustment cost parameter in almost decreased by half ($\kappa_p = 120$). However, shutting off price stickiness (thus allowing firms to costlessly adjust prices) increase inflation volatility, triggers the central bank's inflation stabilization tool and dampens the inflation channel. This situation eventually leads the smoothing channel to dominate since, as one can notice in panel F, yields slightly fall following a policy uncertainty shock.

Apart from the inflation stabilization mechanism, the two other mechanisms of the central bank's Taylor rule—interest smoothing and output stabilization—play little role in explaining the response of yields to the shock (see panels G and H). In fact, Panel G shows that varying the degree of central bank's response to output growth does not lead to substantial changes in the way agents react to the shock. Furthermore, the interest rate smoothing mechanism affects the reaction of yields to the shock only if the central bank commits to an unrealistically large value of the smoothing parameter ($\rho_r = 0.7$ in panel H). The parameter driving the persistence of technology shocks in firms' produc-

tion function, ρ_a, does not affect the responses of yields to policy risk shocks either (see panel C).
The remaining panels (I, J, K and L) show the influence of the persistence of level shocks to fiscal instruments. Only the persistence of level shocks to the capital rate plays a significantly role; this is not surprising given the results reported Figure 1.5 according to which fiscal policy uncertainty was almost entirely captured by uncertainty regarding capital tax policies. Panel J shows that increased persistence of level tax shocks increase the effect that policy uncertainty has on yields; this could be explained by the fact that agents anticipate that if policy authorities were to increase the range within which capital tax rates vary, they would do so do for longer periods of time given the persistence of level shocks.

1.5 Policy Uncertainty and Yields: Empirical Evidence

Given that the above results are based on a calibrated theoretical model, it is straightforward to wonder if they are in line with empirical facts. This section provides empirical evidence that supports my findings relative to the effects of policy uncertainty on yields in the theoretical model. I measure policy uncertainty using data on professional forecasters' disagreement about future values of variables that are in direct link with monetary and fiscal policy actions. I use these data, along with yield curve data and macroeconomic time series, in a structural vector autoregressive (VAR) model.

1.5.1 Data and Methodology

Data

The most important data needed for the analyses are yield curve and policy uncertainty time series. Yield curve data are the 3-month and 6-month treasury bill rates available in the Federal Reserve Bank of St. Louis' FRED database, and the 1-Year, 2-Year, 3-Year, 4-Year and 5-Year discount bond yields constructed by Fama and Bliss, and available at the Center for Research in Security Prices (CRSP).
I use data from three different sources to construct six measures of policy uncertainty (three for monetary policy uncertainty and three for fiscal policy uncertainty); and show that the results of my empirical model are robust to the choice of any of these measures. The first source is the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters; I use their data to construct the two "main" measures of policy uncertainty: one for monetary policy uncertainty, and another for fiscal policy uncertainty. A third measure which is obtained by aggregating the first two, is constructed for the purpose of comparison with other existing measures. The monetary policy uncertainty index is an equally weighted average of forecasters' disagreement on future values of the consumer price index (CPI) and the short-term interest rate (3-Month T-Bill rate). I use one-year

ahead forecasts[19]. Disagreement among forecasters is measured by the cross-sectional coefficient of variation of their point forecasts. The reason for choosing these two forecast variables is almost obvious: both reflect forecasters' perception of future central bank's move regarding monetary policy. Considering the coefficient of variation instead of the standard deviation or the interquartile range allows to reduce the influence of the level of variables on the forecast dispersion—and therefore remain consistent with our definition of a policy uncertainty shock as a mean-preserving spread to shock distributions of future policy variables.

Similarly, the fiscal policy uncertainty index is an equally weighted average of forecasters' disagreement on future values of the real federal government consumption expenditures and gross investment, and the real state and local government consumption expenditures and gross investment. These forecast variables, as stressed by Baker et al. (2013), are also directly influenced by fiscal policy. Again, the cross-sectional coefficient of variation of point forecasts is used to measure disagreement. Finally, the aggregate measure of economic policy uncertainty is an equally weighted average of the monetary and fiscal policy uncertainty indexes[20].

I obtain the second set of data on monetary and fiscal policy uncertainty from Baker et al. (2013). Their policy-related uncertainty index (comparable to the aggregate index presented above) is based three components[21]: (1) the frequency of newspapers' coverage of monetary and fiscal policy uncertainty; (2) an index drawn on tax code expiration data; and (3) economic forecasters disagreement. This latter component is somewhat similar to the index that I construct, except that I use cross-sectional coefficients of variation and one-year forecasts of both CPI and 3-month T-bills to construct the monetary policy uncertainty index whereas Baker et al. (2013) use interquartile ranges and just CPI forecasts. Baker et al. (2013) also construct categorical indexes using similar components as the above; two of their categorical indexes that I also use for in my robustness analyses are their measures of monetary and fiscal policy uncertainty.

The third alternative source of data that I use is Consensus Economic's survey of forecasters. I follow the same strategy as with the Philadelphia FED's data to construct a measure of monetary policy uncertainty using one-year ahead forecasts of the CPI and the 3-month T-bill rate. Fiscal policy uncertainty, however, is captured by forecasters' disagreement on the federal government's budget balance. The main difference between the two sources—Consensus Economics and Philadelphia Fed—is that the latter uses a

[19]The two measures complement each other in the sense that they not only allow to capture monetary policy uncertainty, but they provide different signals about the timing of the actions of a central bank whose objectives include insuring inflation stability. For instance, disagreeing on the value of inflation in one year may signal divergence in forecasters' assessment of the central bank's move in the near future given the delay in the transmission mechanism of monetary policy. In contrast, disagreeing on the value of the 3-month T-bill rate in one year may signal uncertainty about the central bank's decision much later.

[20]It is worth noting that all the averages are computed after normalizing the sub-indices by the value of 1999Q1. 1999Q1 is the start date of an alternative data set on policy uncertainty, namely a data set obtained from Consensus Economics, that I use in my robustness analyses. However, choosing an alternative date would not change the results at all.

[21]For a full description, visit www.policyuncertainty.com.

larger number of forecasters most of the time and has a very large number of forecast variables compared to the former.

I also use output and inflation time series in the VAR model. The output is measured by the log of per capita real output in the US non-farm business sector obtained by dividing real output from the Bureau of Economic Analysis (BEA) by the civilian non-institutionalized population aged 16 and over from the Bureau of Labor Statistics (BLS). Inflation is defined as the percentage change in the CPI for all urban consumers using data available in the FRED database.

Two additional time series capturing short- and medium-term economic expectations (or perception of economic uncertainty) are also used for robustness. The first one, the Chicago Board Options Exchange (CBOE) VIX index, is obtained from Bloomberg; it is usually seen as a short-term (30-day) predictor of market volatility. The second time series, the consumer confidence index, is obtained from the University of Michigan's Consumer Survey. This index is meant to capture expected economic conditions in the medium-term, mostly within five years.

The VIX spans the period 1986M01-2014M12 and Consensus Economics forecast data span the period 1999M01-2014M12. For Bond yields and consumer confidence data, I use the time period 1981M07-2014M12 because data on the main policy uncertainty measures (Philadelphia Fed's data) are available at quarterly frequencies only from 1981Q3. For the same reason, Output and inflation data can only be used from 1981Q3. All monthly data are converted to quarterly frequencies by averaging values of considered months.

Figure A.5 in Appendix A.2 shows the measures of policy uncertainty from the three sources presented above. The figure shows that Baker et al. (2013)'s measures and those based on the Philadelphia FED's data are more volatile than the measures of policy uncertainty based on Consensus Economics' data; this may be because of the small number of forecasters surveyed by the latter. One common feature depicted by the first row of the figure, however, is that policy uncertainty was high in the 1980's and from the late 2000's onward. A further look at the different components of policy uncertainty shows that this pattern was mostly driven by fiscal policy uncertainty, and by monetary policy uncertainty since 2008. A striking observation is that the fiscal policy uncertainty index based on Consensus Economics' data remains low and stable over the entire sample; this is puzzling although the index is based on a completely different forecast variable than the indexes of other sources (budget balance).

The correlations between the measures based on the Philadelphia Fed's data and those by Baker et al. (2013) are small when considered over the entire sample (1985Q1-2014Q4); these correlations are 0.30, -0.13 and 0.17 respectively for economic, monetary and fiscal policy uncertainty measures. The negative correlation for monetary policy uncertainty measures become positive (0.14) when the standard deviation is used to construct the indicator instead of the coefficient of variation. The correlations between the measures of economic and fiscal policy uncertainty are much higher when considered for the sub-sample 1999Q1-2014Q4: they are respectively 0.47 and 0.42. However, the correlation between the measures of monetary policy uncertainty is still small and negative on this sub-sample (-0.10). The correlations between the measures of economic and monetary policy uncertainty based on Fed's data and those based on Consensus Economics' data are strong (between 0.78 and 0.88 respectively). The correlation is nil between the measures of fiscal policy uncertainty from these two sources; which is not surprising given that they are based on completely different forecast variables and given the stability of the index based on Consensus Economics' data compared to other measures, as mentioned above. In spite of the mitigated degree of correlation between the indexes, the results below are robust to the choice of the index to be considered in the VAR.

Methodology

In the benchmark specification, I use a VAR in policy uncertainty, output, inflation, and yields. I focus on the effects of monetary and fiscal policy uncertainty shocks and assume that the effect of a joint shock similar to the one implemented in the theoretical model can be approximated using the sum of the individual effects[22]. Doing so yields results that are more comparable to my theoretical results than looking at the responses of yields to my aggregate economic policy uncertainty index. My models are, therefore, four-variable VARs where the first variable is one of the two specific measures of policy uncertainty (monetary or fiscal), and the last variable is yields for a given maturity. Output is used for identification purposes. Inflation is used because of its important role in the transmission mechanism of policy uncertainty, as shown in the theoretical model. I abstract from including the consumption time series in the specification to reduce the number of estimated parameters given the small sample size; unreported analyses show that its empirical responses to policy uncertainty shocks are similar to that of output[23]. In all specifications and sub-samples considered, the Schwartz Information Criterion (BIC) suggests one lag. The VAR includes a constant and a trend, although the exposition below does not show the trend for simplicity.

I identify policy uncertainty shocks using the sign restriction approach proposed by Mountford & Uhlig (2009) and Uhlig (2005). For this purpose, I restrict the responses of output to be negative on impact and for seven periods following a surprise increase in policy uncertainty. This is to be in line with the findings of a large number of (both theoretical and empirical) papers[24]. The sign restriction approach was originally proposed by Uhlig (2005). It is an agnostic approach as, instead of imposing identifying restrictions on the values of the VAR parameters, it imposes sign restrictions and leave the model determine a range of possible values. Uhlig (2005) and the subsequent literature report the results based on all these possible values—for example by using the median impulse

[22]In a structural VAR where both shocks are orthogonal and identified in the same system, the effect of a joint shock can be approximated by summing up the individual effects.

[23]The correlation between output and consumption, measured by the log of real per capita nondurables and services, is 0.997 in the sample

[24]See among others Fernández-Villaverde et al. (2011), Baker et al. (2013), Mumtaz & Zanetti (2013) and Born & Pfeifer (2014)

response or the percentile bands. This approach also has the advantage to allow to focus only on the identification of the shock(s) of interest. Standard identifying restriction methods in an m-variable VAR require to specify $m(m-1)/2$ identifying restrictions. If one is interested in only $n < m$ structural shocks, there is a priori no reason to identify the remaining $m-n$ (see Uhlig (2005) and references therein), there avoiding the risk of taking an erroneous stand about them . I follow this approach by imposing the restrictions that are necessary to identify the only shock of interest in each of the systems that I estimate.

Let $Y_t,\ t = 1 \cdots T$, denote an n-dimensional vector of variables. The reduced-form VAR model is given by

$$Y_t = c + \sum_{p=1}^{P} B_p Y_{t-p} + u_t, \tag{1.29}$$

where c is an $(n \times 1)$ vector of intercepts, B_p is the matrix of the pth order autoregressive coefficients, P is the optimal lag length, and u_t is the vector of white noise error terms (reduced-form shocks) with variance-covariance matrix Σ.

Denote the vector of structural shocks by ϵ_t. In order to identify these shocks, I adopt the common assumption that there exists a linear mapping between u_t and ϵ_t: $u_t = A_0\epsilon_t$, where A_0 is the impact matrix. The space of possible impact matrices can be characterized by all matrices G of the form $G = \tilde{A}_0 D$ where D is an orthogonal matrix ($DD^{'} = I$), and $\tilde{A}_0$ is an arbitrary orthogonalization of Σ.

Assume for simplicity that $\tilde{A}_0$ is the Cholesky decomposition of Σ and that the policy uncertainty variable is ordered first in Y_t. The first column of G represents the immediate impact, or impulse vector, of a one standard error innovation to the policy uncertainty shock ($\epsilon_{1,t}$). I identify $\epsilon_{1,t}$ by selecting the first column of an orthogonal matrix D which satisfies my restriction that output does not increase on impact and during the first seven quarters after a sudden increase in (monetary or fiscal) policy uncertainty. Only the impulse vector to this shock needs to be characterized, which I do below; the reader is referred to Mountford & Uhlig (2009) for more details on this identification strategy.

Let $g_1 = \tilde{A}_0 d_1$ be a generic impulse vector for the shock of interest, where d_1 is the first column of an orthogonal matrix D. Denote by $r_{j,i}(h)$ the impulse response of the jth variable to the ith column of $\tilde{A}_0$ at horizon h (that is, the reduced-form IRF), and by $r_i(h)$ the n-dimensional column vector $[r_{1,i}(h), \cdots, r_{n,i}(h)]$. The n-dimensional impulse response $r_{g_1}(h)$ to the impulse vector g_1 is given by

$$r_{g_1}(h) = \sum_{i=1}^{n} d_{i,1} r_i(h), \tag{1.30}$$

where $d_{i,1}$ is the ith entry of d_1.

Following Mountford & Uhlig (2009) approach, I select the impulse vector d_1 that solves the minimization problem:

$$d_1^{\star} = \underset{d_1^{'} d_1 = 1}{\operatorname{argmin}}\, \Psi\left(\tilde{A}_0 d_1\right), \tag{1.31}$$

where the criterion function $\Psi\left(\tilde{A}_0 d_1\right)$ is given by

$$\Psi\left(\tilde{A}_0 d_1\right) = \sum_{l=0}^{7} f\left(\frac{r_{y,g_1}(l)}{s_y}\right) \tag{1.32}$$

and the loss function, f, is such that $f(x) = 100x$ if $x > 0$ and $f(x) = x$ if $x \leq 0$; s_y is the standard deviation of the reduced-form innovation to output. The criterion $\Psi\left(\tilde{A}_0 d_1\right)$ therefore strongly penalizes the impulse vectors that generate a positive output response.

1.5.2 VAR Results

In this section I present the estimated impulse responses to the identified policy uncertainty shocks. I first present the results of the benchmark specification (four-variable system including policy uncertainty measures, output, inflation and yield). Next, I present the results from an extended specification (benchmark specification augmented with the VIX and the consumer confidence index) and the results from various robustness checks including alternative measures of policy uncertainty and different sub-samples. Under each specification, I estimate two systems: one with the main measure of monetary policy uncertainty, and the other with the main measure of fiscal policy uncertainty. In the subsequent figures the lines depict the median response to a one-standard deviation increase in the log of policy uncertainty, from 2000 replications of the reduced-form VAR using Kilian (1998)'s bias-corrected bootstrap procedure. The lower and the upper bounds of the confidence bands represent the 2.5th and the 97.5th percentiles respectively; and the vertical shaded band at the beginning of output's response represents the horizons where the response of output to the shock is restricted to negative values. Except for yields, all the variables in the figures are logged.

Benchmark Specification

Figure 1.8 shows the impulse responses to the identified (monetary and fiscal) policy uncertainty shocks. The purple dotted lines depict the effect of the monetary shock while the blue dashed lines show the effect of the fiscal shock. The sign restriction on the response of output is mostly important for the initial effect as, without this restriction, the shock leads to an impact increase in GDP. A one standard deviation increase in monetary (resp. fiscal) policy uncertainty corresponds to a 16 (resp. 25) percentage points increase. The figure shows that the fiscal policy uncertainty shock is short-lived. Indeed, a shock to fiscal policy uncertainty is followed by a rapid decline of this variable to its pre-shock value within almost one year after the shock. The monetary policy uncertainty shock is more persistent than the fiscal shock; the monetary policy uncertainty index is almost back to its pre-shock value after ten quarters.

Consistently with the results of the theoretical model, a sudden increase in (monetary or fiscal) policy uncertainty leads to an impact increase in interest rates of all maturities. The shock triggers an immediate increase in yields, followed by a gradual decline. The effect of

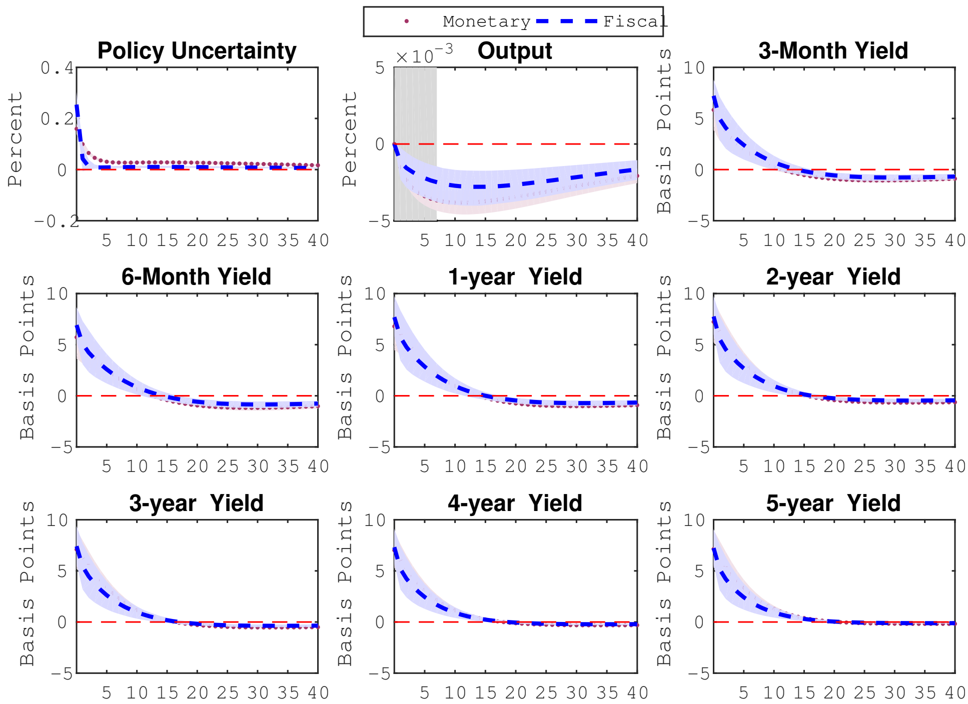

Figure 1.8: Impulse responses to one-standard deviation policy uncertainty shocks: Benchmark Specification (1981Q3-2014Q4).

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

the shock on yields turn negative for all maturities some 13 to 22 quarters after the shock, with the timing increasing with maturity. Yields, then, go back to their pre-shock levels. The pattern followed by the responses of yields is, therefore, very qualitatively similar to the pattern of theoretical impulse responses although the latter turn negative more quickly. A one percent increase in fiscal policy uncertainty leads to a rise in the 3-month T-bill rate by 28 (= 7/0.25) basis points on impact; this is followed by a gradual decline to a minimum of -3.1 basis points after seven years before returning to the pre-shock level. The 5-year interest rate jumps up by 29 basis points on impact before gradually declining as well. Yields of remaining maturities follow a similar pattern.

Figure 1.8 also shows that following a one percent increase in monetary policy uncertainty, the 3-month T-bill rate rises by 37 (= 6/0.16) basis points on impact before quickly declining; the effect of the shock becomes negative after almost fifteen quarters and reaches a minimum of -6.6 basis points after seven years before returning to its pre-shock level. The 5-year rate also reacts to the shock, increasing by 45 basis points on impact before gradually declining, reaching a minimum of -1.1 basis points nine years later and returning to its pre-shock level. The remaining interest rates react similarly.

Extended Specification

Policy uncertainty is influenced by contemporaneous economic developments and expectations about the future, and forecasters perceive signals about future economic conditions that they can account for when making their forecast for policy-related variables. Consequently, it is straightforward to suppose that the measures of policy uncertainty constructed above may, as well, capture economic uncertainty in general, instead of just policy uncertainty. For this reason, I include two widely used measures of economic uncertainty, namely the VIX and the consumer confidence index, in an alternative specification of the VAR in order to improve the identification of policy uncertainty shocks. The systems estimated under this extended specification are, therefore, seven-variable VAR models; once again, two systems are estimated: one with the measure of fiscal policy uncertainty, and the other with the measure of monetary policy uncertainty. Figure 1.9 plots the responses of yields to the monetary and fiscal policy uncertainty shocks. As previously, the purple dotted lines depict the effects of the monetary shock and blue dashed lines depict those of the fiscal shock.

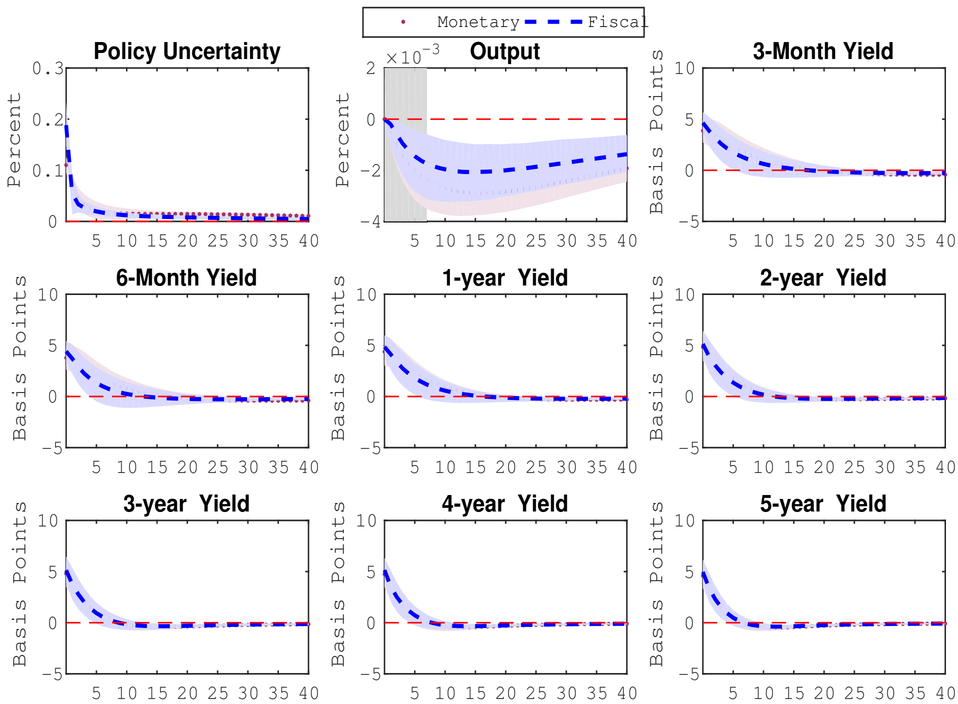

Figure 1.9: Impulse responses to one-standard deviation policy uncertainty shocks: Extended Specification (1986Q1-2014Q4).

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

The figure shows that the responses of yields on impact and in the quarters following the shocks are qualitatively similar to what we observed in smaller systems. The shocks lead to an impact increase in interest rates, followed by a relatively gradual decline; as under the benchmark specification, the effects turn negative before the yields go back to their pre-shock levels. The impact effects are, however, slightly smaller than in the four-variable VAR. Indeed, a one percent increase in fiscal (resp. monetary) policy uncertainty leads the 3-month T-bill rate to increase by 25 (resp. 36) basis points compared to 28 (resp. 37) basis points under the benchmark. The impact effect on the 5-year interest rate also decreases by 3 and 2 basis points, respectively for the fiscal and the monetary shocks.

Alternative Measures of Policy Uncertainty and Sensitivity

In order to further assess the robustness of the results presented above, I replace the main measures of policy uncertainty used in the benchmark specification with the alternative measures presented in Section 1.5.1, namely (1) the measures of Baker et al. (2013) mostly based on the share of newspaper articles mentioning economic policy uncertainty, and (2) the measure of forecasters disagreement based on Consensus Economics' data. I also estimate the benchmark model using different lag lengths and on different sub-samples. The results of these additional analyses are presented in Appendix A.2.

Figures A.6 and A.7 show that with the alternative measures of policy uncertainty, yield responses to the fiscal policy uncertainty shock are qualitatively similar to benchmark results. This is also the case when Baker et al. (2013)'s measure of monetary policy uncertainty is used. This means that the shocks trigger and impact increase in yields followed by a gradual decrease; eventually the effects turn negative for a few quarters before vanishing. It is worth noting, however, that Baker et al. (2013)'s data overestimate the effects of the shocks compared to the benchmark while Consensus Economics's data tend to underestimate them. A somewhat surprising result is the responses of short yields to the monetary policy uncertainty shock when Consensus Economics' data are used to measure the monetary policy uncertainty. As we can see, the impact effects of the shocks on yields of maturities lower than 4-year is nil; only for four- and five-year yields does the shock have a positive impact effect.

Figures A.8 and A.9 show that the yields respond to (both monetary and fiscal) policy uncertainty shocks positively even when different sub-samples of the data are considered. The shocks appear to have higher effects in the first half of the sample (1981Q3-1998Q4) than in the second half (1999Q1-2014Q4). Note that the responses of yields to the two shocks turn negative at some point only in the first half of the sample. The benchmark system is also estimated with more lags in the reduced-form VAR and the responses of yields to the shocks remain qualitatively the same, although less smooth. Figure A.10 shows this in a system with three lags.

The results of the VAR analyses presented in this Section 1.5 provide strong evidence the data support the results of my theoretical model that policy risk shocks trigger an increase in yields, and that this is due to an inflation risk premium that outweighs the smoothing effect.

1.6 Conclusion and Policy Implications

I have shown that when the economy is subject to unpredictable policy risk shocks, the level of the median yield curve is lower, its slope increases and term premiums decrease on average. Most of these effects are attributable to fiscal policy uncertainty which, in turn, is driven by capital tax policy uncertainty. The overall long-term downward effect on yields and premiums is due to the asymmetric impact of positive and negative shocks to the volatility of policy instruments, negative shocks having more amplified effects compared to positive shocks of the same magnitude. Impulse response analyses show that following a policy risk shock, yields of all maturities increase. This is because the fall in yields triggered by higher demand for bonds by households, in order to hedge against expected consumption volatility in the aftermath of the shock, is outweighed by the increase in yields due to higher inflation term premiums. The most important feature driving the observed response of yields to policy risk shocks is investment adjustment costs as they determine the pace of resource reallocation. These costs prevent households from lowering investment in order to reduce the distortions created by capital tax policy risks.

Empirical analyses show that the findings from the theoretical model are supported by empirical evidence. I construct indexes of monetary and fiscal policy uncertainty using data on forecasters' disagreement about future values of policy-related variables, and identify policy uncertainty shocks in a structural VAR model using sign restrictions. An increase in (monetary or fiscal) policy uncertainty leads interest rates of all maturities to jump up on impact before returning to their pre-shock values, consistently with the theoretical impulse responses.

Recent developments in the "policy uncertainty" literature show that policy risk shocks can have sizable and persistent negative real effects. This paper complements these findings by stating that the effects on the financial side of the economy can also be important. My findings are relevant for both fiscal and monetary policy authorities. First, higher policy uncertainty can be problematic for government finances. My analyses show that the long-run effect of equally likely positive and negative policy uncertainty shocks is a decrease in interest rates on government bonds. Given that in reality higher policy uncertainty tends to materialize more often than higher policy transparency, especially as far as fiscal policy uncertainty is concerned, this result implies that policy uncertainty can be a threat for long-term fiscal sustainability, or at least permanently increase the cost of government debt. Second, higher policy risk can also hamper the effectiveness of monetary policy in a context of economic slowdown. The results on impulse response functions suggest that since yields mostly react to fiscal policy risk shocks, greater government indecisiveness regarding its policy can, during an economic recession already exacerbated by policy uncertainty itself, undermine the transmission mechanism of monetary policy by holding interest rates high after a monetary easing; this would considerably slowdown the pace of economic recovery. My results, thus, call for more policy transparency and more coordination between monetary and fiscal authorities.

Chapter 2

Financial Markets Convergence and Determinants of Risk Premium Differentials

2.1 Introduction

The rapid integration of global financial markets over the last two and a half decades, largely characterized by increased access of Emerging Market Economies (EMEs) to world capital markets, raises questions about the extent to which government bond yields and risks in EMEs have converged towards the levels found in Advanced Economies (AEs). The questions are even more relevant given the fact that available data suggests that the share of foreign holdings of local currency government bonds in EMEs has more than tripled between 2004 and 2014 (see the solid blue line in Figure 2.1) shows. Indeed in five out of eight EMEs studied in this paper, this share grew from 10 percent in 2004 to 30 percent in 2014, as the red dotted line in Figure 2.1[1]. Moreover, the investor base has considerably increased during the last decade (see Arslanalp & Tsuda (2014)). This paper looks at the market for government bonds in 12 advanced economies and 8 emerging market economies, during the period 1999-2012, and considers the particular question of whether or not there has been any convergence of risks between EMEs and AEs. A distinction is made between default risk and other types of risk (such as inflation, liquidity and exchange rate risk), and the focus of the paper is on the latter.

The literature on risk convergence between AEs and EMEs is substantial and has continued to grow during the last few years. One recurring observation in this literature is the rapid decrease in government bond interest rate spreads between AEs and EMEs, mainly pointing to convergence in credit risks[2]. What drives this convergence remains controversial however. While one strand of the literature attributes the contraction of Emerging Market spreads to global factors such as increased investors' appetite for risk and easy global macroeconomic conditions (see González-Rozada & Levy-Yayeti (2008) and Csonto & Ivaschenko (2013) among others), another attributes it to improvements in local fundamentals (see for instance Eichengreen & Mody (1998), Baldacci et al. (2008), Arbatli (2011)). Some authors have argued that both factors played non-negligible roles in explaining the spread compressions that we saw during the last few decades (Calvo et al. (1996) and Hartelius et al. (2008)).

The exact contribution of each set of factors to tightening the spread is hard to disentangle. This tightening occurred when EMEs were considerably improving their economic fundamentals, reducing fiscal strains and debt burdens, transitioning to more market-based institutions, and implementing various structural reforms (Calvo et al. (1996), Hartelius et al. (2008) and Ciarlone et al. (2009)). At the same time, systemic AEs went through many periods of very low and stable interest rates, and investors were seeking alternative sources of high yields (Calvo et al. (1996) and Hartelius et al. (2008)). Examples of such periods include the many years of excess liquidity conditions that were triggered by the

[1] Data on the share of foreign holdings of government bonds denominated in local currencies are available for only five out of eight emerging market economies studied in this paper. Those five countries are: Hungary, Malaysia, Mexico, The Philippines and South Africa.

[2] Some papers using or referring to interest rate spreads as measures of credit/default risk are Duffie et al. (2003), Darrell & Singleton (2003), Diana & Gemmill (2006), Eli M. et al. (2007), Baldacci et al. (2008) and Blot et al. (2016)

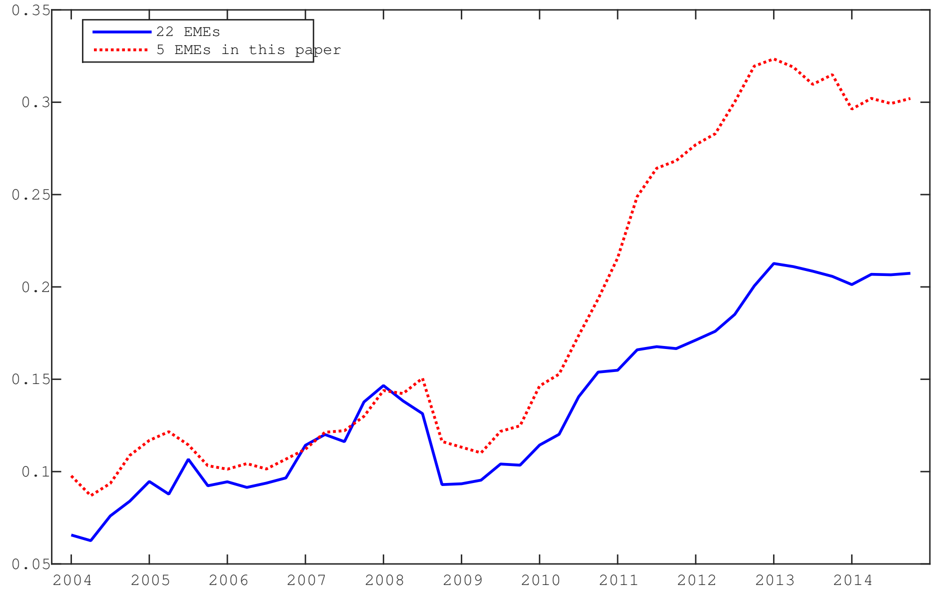

Figure 2.1: Share of foreign holdings of local currency denominated government bonds in Emerging Market Economies (EMEs).

Notes: The blue solid line represents the share for 22 EMEs (Argentina, Brazil, Bulgaria, China, Colombia, Egypt, Hungary, India, Indonesia, Latvia, Lithuania, Malaysia, Mexico, Peru, The Philippines, Poland, Romania, Russia, South Africa, Thailand, Turkey, Ukraine), and the red dotted line represents the share for 5 of the 8 EMEs studied in this paper (Hungary, Malaysia, Mexico, The Philippines, South Africa), and for which the data are available. Quarterly data from 2004 to 2014. Source: Arslanalp & Tsuda (2014).

responses of advanced economies to the bursting of the equity bubble in the early 2000s and the global financial crisis of the late 2000s.

In its analysis of convergence between EMEs and AEs, the literature has so far put less emphasis on studying risks other than credit risks. This paper sheds more light on this undeveloped strand of the literature by studying these other sources of risk (such as inflation, liquidity, interest rate, exchange rate risks, etc.). Indeed, losing money is a primary concern for investors and credit risk is directly related to loss of money as is it the risk that the borrower will not be able to (fully) repay its debt; however, non-credit sources of risk might as well be very important. This paper seeks to deepen the analysis of those other sources of risk. One of the motivations here is that while these other sources of risk might be small relative to default risk, they are nonetheless taken into account by investors in their asset pricing decisions. For instance, all things equal, in periods of high capital mobility, countries with lower risks associated with factors other than default are expected to attract more inflows or less outflows. This is beneficial for the country, say, by making it less vulnerable to sudden stops. Provided that there is convergence of credit risk, continuous and important capital movements between EMEs and AEs during periods of "normal" global macroeconomic and liquidity conditions would suggest that the two types of economies differ in their levels on this second category of risks.

Besides tackling an underdeveloped aspect of studies on "bond risk convergence", this paper contributes to the literature in three ways. First, it shows that forward risk premium differentials can be used to measure differences between EMEs and AES in risks other than default risks. Indeed, I use a simple two-period portfolio choice model to make the theoretical case that the forward risk premium differential between two countries captures all the risk premiums considered by bond buyers, except the credit risk premiums which is often measured by spreads. The second contribution of the paper is to use the constructed measure of risk in order to show that the levels of risk other than default in EMEs have not converged towards those seen in AEs, between 1999 and 2012. Finally, the paper seeks to understand the reasons behind this absence of convergence, by investigating the determinants of risk differentials between the two categories of country.

Investigating the differential in risk premiums between EMEs and AEs provides useful information about the behavior of investors when global financial conditions change. It essentially helps in the analysis of the extent of capital outflows from EMEs in the current period of tightening financial conditions. Suppose for instance that the differential has decreased and EMEs risks have converged towards AEs' levels; then, capital reversals—i.e. outflows of funds that investors transferred to EMEs while seeking better rewards during the long period of loose financial conditions—will most likely be limited following the normalization of monetary policies in AEs. In contrast, if these other sources of risk have remained high in EMEs relative to their AEs counterparts, tighter monetary conditions in AEs will lead to larger capital outflows from emerging countries, potentially creating some economic damage. Furthermore, analyzing the determinants of the differential helps to inform policymakers in EMEs about whether their efforts to deepen local financial markets have led to conclusive results through the reduction of all kinds of risks. The absence of convergence will therefore mean that whenever global financial conditions shift, these economies need to take additional measures in order to face potentially damaging economic volatility.

Using three complementary statistical and econometric tools (namely correlations, principal component analysis, and maximum differentials), I show that non-credit risk premiums in EMEs have not converged towards those of AEs over the period 1999-2012. Indeed, the analyses show that risk premiums among AEs economies are positive and very high while they remain very low and most of the time negative between AEs and EMEs. Moreover, the principal component analysis shows that there is not a common factor that drive risks premiums from these two categories of country. The previous two exercises are complemented by a third one that consists in constructing artificial time series aiming at capturing the largest risk spread for each period. The evolution of these time series does not show a clear decreasing pattern and their regression on a time trend shows negative but non-significant slopes. These results on the absence of risk convergence are robust across different maturities of government debt instruments (3-month, 6-month, and 5-year).

The investigations on the determinants of risk premium differentials are done through a panel regression. The results of the analyses lead to the conclusion that countries differences in their local economic fundamentals, and in their degree of political risks are the main drivers of risk premium differentials. In particular, economic growth and the liquidity of local financial markets are found to be the main macroeconomic fundamentals whose differences between AEs and EMEs drive risk premium differentials. Inflation differentials also play an important role on short-term risk premium differentials. The results also show that investors highly value strong economic stability in periods of high market volatility and economic crisis. Investigations region-specific differences in investors' asset pricing behavior reveal that Asian and European EMEs tend to have higher risk premiums.

In the next section, I show how risk premiums are computed and present a simple theoretical model which shows that forward risk premium differentials can be used to capture risk differences in risks other than default between two countries. Section 2.3 presents the data present some descriptive statistics on forward risk premiums. Section 3.3 reports my analyses on the convergence between EMEs and AEs. Section 2.5 presents the the empirical model used to investigate the determinants of risk premium differentials and exposes the results of these investigations. Section 2.6 concludes and presents some policy implications.

2.2 Risk Premiums and Risk Premium Differentials

2.2.1 Risk Premiums in Interest Rates

In order to assess the risk associated with a given country's government bond, many authors in the "bond risk convergence" literature have used the interest rate spread between the country of interest and a benchmark country, usually the US (see among others Eichengreen & Mody (1998), González-Rozada & Levy-Yayeti (2008), Hartelius et al. (2008), Ebner (2009), Csonto & Ivaschenko (2013), Kennedy & Palerm (2014)). The most commonly used measures are the Emerging Market Bond Index (*EMBI*) and the *EMBI Global* index. The fact that most papers in the literature refer to the interest spreads as a measure of the risk of default (see Duffie et al. (2003), Darrell & Singleton (2003), Diana & Gemmill (2006), Eli M. et al. (2007), Schuknecht et al. (2009), Bernoth et al. (2012) and Blot et al. (2016) among many others) reflects the belief that a decline in the value of this variable is associated with a decline in the probability of default perceived by investors at a given point in time—due to the information available to them on potential improvements in the country's fundamentals, global conditions, or other factors. Changes in the risk of default can explain cross-border movements of funds between countries, but it is also obvious that risk factors other than default play a non-negligible role.

The objective of this section is to look at sources of risk other than default; this has less been the focus of empirical research on risk convergence. The paper aims to analyze the convergence of emerging bond markets in terms of those risks not associated with the probability of default (credit risk). The strategy is not to take each source individually but, instead, to construct an aggregate measure that encompasses all (or most) of these other kinds of risk. In order to achieve this goal, I use forward risk premiums on government zero-coupon monthly yields of maturities 3, 6, and 60 months. In the remainder of this section, I show that forward risk premium differentials can be used as an aggregate measure of risks other than default.

Denote the spot rate of a $k-$period government zero-coupon bond at time t by $r_t(k)$. The *$n-$period-ahead $k-$period forward rate* is defined by

$$f_t(n,k) = \frac{1+r_t(n+k)}{1+r_t(n)} - 1 \qquad for \quad n,\ k = 1, 2, \ldots \tag{2.1}$$

This is the interest rate an investor would require today to hold an asset during k periods in the future beginning n periods from now. To be more explicit, consider a contract made at time t, which gives the investor the right to hold some amount of k-period bonds that the government will issue at time $t+n$. The $n-$period-ahead $k-$period forward rate is the rate at which the investor will accept the contract. From the investor's perspective such a forward contract is risky given that the future state of the economy is unknown, especially between the time when he signs the asset and the time that he actually holds the bonds. He will therefore require a higher compensation than the expected spot rate of a k-period bond issued in period $t+n$. I therefore define the *$n-$period-ahead $k-$period forward risk premium* by

$$fp_t(n,k) = f_t(n,k) - E_t\left(r_{t+n}(k)\right) \tag{2.2}$$

the difference between the forward rate and the expected spot rate. E_t denotes the expectation conditional on the time t information set. Defined this way, risk premiums on government assets can be estimated for every country, which contrasts with spreads that are excess returns on a bond relative to a "riskless" counterpart. This definition of risk premium therefore allows for the study of the evolution of risk even for countries usually considered as benchmark. Moreover, the sign convention means that the risk premium is positive when the forward rate exceeds the expected future spot rate. This measure of risk premium is largely used in the macroeconomic literature; indeed, it was used by Backus et al. (1989) and Gürkaynak & Wright (2012) to study risk premiums in the U.S. term structure and the expectations hypothesis. However, it is used here for a different purpose, namely for the analysis of convergence in risks other that credit risk.

Now, I present a simple theoretical framework that determines how a portfolio manager derives the forward risk premium differential between countries' government bonds, and more importantly, that shows that this differential captures all potential sources of risk except credit risk. This result of the theoretical model will motivate the specification of the empirical model estimated in Section 2.5.

2.2.2 A Simple Portfolio Choice Model

The model presented here is a two-period version of the mean-variance model of portfolio choice proposed by Bernoth et al. (2012) and also used by Schuknecht et al. (2009) to study the convergence of sovereign risk premiums in the EMU; most of the notation here is therefore similar to theirs. Considerable changes are brought to the model however, and the focus is on forward risk premiums instead of spreads. The details of the derivations are presented in Appendix B.1, only a summary of the model is presented below.

Consider two governments, one an emerging market economy and the other an advanced economy; we will take the former as the domestic economy. Each government issues an one-period maturity government bond that pays interest r_t for the domestic economic and $r_t^\star$ for the foreign economy. Each government also issues a forward contract which, if held by an investor in the first period, promises to pay interest f_t on bonds issued in the second period in the domestic economy, and $f_t^\star$ when issued by the foreign government.[3]

Before getting into the optimization problem, let's first present an intuition of what a forward risk premium differential captures. Imagine two investment strategies: one that consists in buying a forward contract today, and the second that consists in waiting until tomorrow and buying the one-period government bond. Although these two investment strategies are based on the government bonds that will be issued tomorrow, they differ in the risk they bear. Indeed, in addition to the risk factors considered by an investor who chooses the second strategy, another investor who chooses the first strategy will take into account potential risky events that could materialize between now and tomorrow in its asset pricing. This additional risk takes into account every potentially risky event except default since the government cannot default on an asset that is not yet issued. What I call *forward risk premium differential* is the difference in terms of this additional risk between two countries. The model below aims at formalizing this intuition using some restrictive assumptions to guarantee tractability.

Now, consider a domestic risk-averse portfolio manager choosing among these four securities in the first period and two securities in the second[4]. Each period, he faces costs l that reduce his wealth and that could be attributed to poor economic and institutional characteristics of the domestic country. The losses due to poor economic characteristics derive from low liquidity of the local financial market, high reinvestment risk, high inflation, high taxes triggered by bad fiscal performance (and exchange rate fluctuations if the assets are denominated in different currencies as will be the case in our empirical analyses presented later), etc. These costs are proportional to the amount of domestic transactions carried out by the portfolio manager. Securities issued by the foreign government are considered as benchmark in the market and the costs related to them are normalized to zero.

The portfolio manager maximizes a utility function that depends positively on the ex-

[3]Because this is a two-period model, forward contracts are not issued in the second period by any government.

[4]In the model description, I do the analysis from the perspective of a domestic investor, but the reasoning and the results are analogous if the analysis is conducted from the perspective of a foreign portfolio manager

pected wealth, $E_t[w_{t+2}]$ and negatively on its variance, $Var_t[w_{t+2}]$:

$$Max\ U\{E_t[w_{t+2}], Var_t[w_{t+2}]\},\ U_1 > 0,\ U_2 < 0. \tag{2.3}$$

The portfolio manager takes all investment decisions in the first period, even decisions regarding the composition of the portfolio of bonds in the second period. This assumption is made to ensure tractability. Practically, one could think about an investor that requires his portfolio manager to provide an investment plan for both periods and asks him to commit to this plan. This assumption, However, comes with some loss of realism for the model since it is known that investors update their decisions intertemporally. Nevertheless, the intuition behind the results of the model holds regardless of the assumption.
The decision on the portfolio composition in period two is, thus, based on the expected values of all information needed by the portfolio manager (for instance expected wealth, rates, costs and probabilities of default). In the first period, the portfolio manager therefore chooses the fractions θ_t^{ds} and θ_t^{fs} of his wealth w_t to allocate to domestic and foreign government bonds respectively; the fractions θ_t^{df} and θ_t^{ff} to allocate to domestic and foreign forward contracts respectively; and the fractions θ_{t+1}^{ds} and $1-\theta_{t+1}^{ds}$ of his expected second period's wealth ($E_t[w_{t+1}]$) to allocate to domestic and foreign government bonds respectively. The portfolio manager faces the following budget constraint:

$$\theta_t^{ds} + \theta_t^{df} + \theta_t^{fs} + \theta_t^{ff} = 1 \tag{2.4}$$

As Bernoth et al. (2012), I assume that domestic securities are subject to the risk of partial default, but foreign assets are risk-free. With a probability of $1 - P(x_t),\ 0 \leq P(x_t) \leq 1$, the domestic government will default on its debt, repaying only a fraction, $\alpha_t \in (0, 1)$, of it. x_t denotes the set of variables that influence this probability. In the following, I will use P_t instead for convenience and denote its period's two expected value by P_{t+1}. The expected wealth and its variance are given by:

$$\begin{aligned} E_t[w_{t+2}] = {} & (1 + E_t r_{t+1})\,\theta_{t+1}^{ds} E_t[w_{t+1}] P_{t+1} + \alpha_{t+1}(1 + E_t r_{t+1})\,\theta_{t+1}^{ds} E_t[w_{t+1}](1 - P_{t+1}) \\ & - \theta_{t+1}^{ds} E_t[w_{t+1}] l_{t+1} + (1 + f_t)\,\theta_t^{df} w_t P_{t+1} + \alpha_{t+1}(1 + f_t)\,\theta_t^{df} w_t (1 - P_{t+1}) \\ & + \left(1 + E_t r_{t+1}^{\star}\right)\left(1 - \theta_{t+1}^{ds}\right) E_t[w_{t+1}] + (1 + f_t^{\star})\, w_t \theta_t^{ff} \end{aligned} \tag{2.5}$$

$$Var_t[w_{t+2}] = (1 - \alpha_{t+1})^2 \left[(1 + E_t r_{t+1})\,\theta_{t+1}^{ds} E_t[w_{t+1}] + (1 + f_t)\,\theta_t^{df} w_t\right]^2 P_{t+1}(1 - P_{t+1}) \tag{2.6}$$

where the expected wealth at the end of the first period, $E_t[w_{t+1}]$, is given by

$$E_t[w_{t+1}] = (1 + r_t)\,\theta_t^{ds} w_t P_t + \alpha_t (1 + r_t)\,\theta_t^{ds} w_t (1 - P_t) - \left(\theta_t^{ds} + \theta_t^{df}\right) w_t l_t + (1 + r_t^{\star})\,\theta_t^{fs} w_t \tag{2.7}$$

Combining the first-order condition with respect to θ_{t+1}^{ds}, with the first-order condition with respect to θ_t^{df} yields the following equilibrium condition:

$$\frac{1+f_t^\star}{1+f_t} - \frac{1+E_t r_{t+1}^\star}{1+E_t r_{t+1}} = \frac{l_{t+1}}{1+E_t r_{t+1}} - \frac{1+E_t r_{t+1}^\star}{1+f_t} l_t \tag{2.8}$$

The last equation can in turn be rearranged to obtain the following

$$(f_t - E_t r_{t+1}) - \left(f_t^\star - E_t r_{t+1}^\star\right) = l_t + \Psi_t \tag{2.9}$$

where $\Psi_t = l_t \left[E_t r_{t+1}^\star + \left(1 - E_t r_{t+1}^\star\right) E_t r_{t+1}\right] - l_{t+1}(1+f_t) + f_t^\star E_t r_{t+1} - f_t E_t r_{t+1}^\star$ is a term that involves various covariances among the different interest rates as well as the expected cost.

There are at least two interesting conclusions about (2.9). First, the forward risk premium differential between the domestic and the foreign government bonds is a function of the factors that reduce investors' wealth and of the covariance between the various rates on the debt instruments. Second, the forward risk premium differential does not depend on the probability of default (neither does it depend on the haircut rate, $1-\alpha_t$), and therefore does not capture credit risks. This result supports our statement that the forward risk premium differential captures differentials between countries in terms of "all" risks, except the credit risk. The model focuses on a one-period ahead one-period forward contract, but generalizing the framework to a $k-$period ahead $n-$period forward contract will not change the main conclusion of the model though the derivations will be more complicated.

In order to further develop the reader's understanding of the intuition behind the above, I breakdown the differential into two components: the forward rate differential ($f_t(n,k) - f_t^\star(n,k)$) and the expected spread ($E_t[r_{t+n}(k) - r_{t+n}^\star(k)]$), and show how each component varies according to potential risk factors. Consider an investor establishing two forward contracts at time t: one issued domestically at the rate f_t and the other issued by a foreign government at the rate $f_t^\star$. The contracts are on k-maturity bonds that will be issued by the respective governments at $t+n$, and therefore come to maturity at $t+n+k$. Let's see how $f_t - f_t^\star$ varies with respect to potential factors affecting risks between t and $t+n+k$. On the one hand, the investor will account for potential risks that could erode the value of his assets between t and $t+n$, this includes changes in the difference in inflation, liquidity, exchange rate, etc. across countries, but not default risks since the bonds will be issued only at $t+n$. On the other hand, the investor will also take into account the same risks as previously between $t+n$ and $t+n+k$ but additionally account for potential risks of default by any of the two governments.

The expected spread between the interest rates on bonds that will be issued by both governments in $t+n$—$E_t[r_{t+n}(k) - r_{t+n}^\star(k)]$—varies according to changes in the gaps between the two countries in terms of all risk factors between $t+n$ and $t+n+k$, i.e. credit risk and risk factors other than default. Since the expectations are all based of information available at t, the risk differential captured by the spread is similar to the one included in the forward rate differential between $t+n$ and $t+n+k$. These cancel in the risk premium differential, which leads to the fact that the risk premium differential varies according to risk factors other than default between t and $t + n$.

2.3 Data and Selected Statistics on Risk Premiums

2.3.1 Data

To study convergence between EMEs and AEs, I construct forward rates using data on government bond yields with 3-, 6-, 12-, 60-, and 120-month maturities. These data are collected on the websites of national central banks and statistical agencies of the countries studied. In order to construct the forward risk premiums, the expected spot rates on government bonds are needed; these data are obtained from Consensus Economics. Consensus Economics conducts surveys every months and asks professional forecasters to provide their forecasts about major economic and financial variables in a number of countries. The forecast variables include interest rates on government bonds. I also use the average of past realizations of the spot rate over the desired horizon as another measure of expected spot rate, in order to conduct some robustness analyses. For instance, the expected rate of a given asset in a one year horizon would be the average of the realized rates during the current month, and during the last eleven months (the results of the convergence analyses based on this measure are presented in Section 2.4.4).

In Section 2.5, macroeconomic fundamentals (inflation, reserves, liquidity, government balance, GDP growth, and real exchange rate) of the countries of interest are used to investigate the determinants of risk premium differentials between EMEs and AEs. They are obtained from the IMF's International Financial Statistics and the Bank for International Settlements (BIS). In the investigations, I also use data that measure the global economic situation, namely global liquidity measured by the Fed funds rate, and the global investors risk appetite measured by the VIX. The former is obtained from the Federal Reserve Bank of St. Louis' FRED database, and the latter form Bloomberg. Data on political risks constructed by the Political Risk Services (PRS) Group are also used in the study.

The panel of countries includes twelve Advanced Economies (Australia, Belgium, Canada, France, Italy, Malta, New Zealand, Norway, Sweden, Switzerland, the United Kingdom, and the United States), and eight Emerging Market Economies (Croatia, Hong Kong, Hungary, Malaysia, Mexico, The Philippines, Singapore, and South Africa). This selection of countries is imposed by the availability of data on government bond yields. The categorization of countries as AEs or EMEs is from IMF's April 2014 World Economic Outlook. The data span the period $1999 - 2012$. Monthly data are used in this section and in Section 3.3 where the convergence analyses are conducted. Quarterly data are used to investigate the determinants of risk premium differentials in Section 2.5.[5] This difference in the frequency is due to the availability of only quarterly data on most of macroeconomic fundamentals. The first point of the sample period corresponds to the first period

[5]Quarterly data on the forward risk premiums are obtained but taking the values of the last month in the quarter. As stated in Section 2.5, this helps in addressing the simultaneity problem in the regressions.

of 1999 ($M01$ for monthly series and $Q1$ for quarterly series), this date has been chosen in order to consider only the Post-EMU (European Economic and Monetary Union) period. The last points of the sample periods are $2012M06$ and $2012Q2$, respectively for monthly and quarterly series; this date has been chosen to avoid the period during which some advanced-economy governments experienced negative nominal interest rates on their securities.

2.3.2 Selected Statistics on Risk Premiums

The definition presented (2.2) above is used to compute the series of forward risk premiums on government securities for three different maturities. Since available data include interest rates on only 3-, 6-, 12-, 60-, and 120-month government bond yields, the only measures of risk premiums that can be computed using (2.2) are the $3-$month-ahead $3-$month forward risk premium, the $6-$month-ahead $6-$month forward risk premium, and the $60-$month-ahead $60-$month forward risk premium. In the remainder of the paper, I will refer to these measures as $3-$, $6-$, and $60-$month risk premiums since there is no ambiguity. In Table 2.1 I report some descriptive statistics for the risk premiums.
The table shows that risk premiums are not very volatile. The coefficient of variation is very small, i.e. risk premiums do not vary much relative to their mean. However, longer maturities are more stable relative to their mean and much more persistent than risk premiums on shorter maturities. Risk premiums also increase with maturity, revealing the upward slope of the yield curve, which is a well-known feature of the term structure of interest rates. These observations are analogous to some stylized facts of the term structure of interest rates' literature. Diebold & Li (2006) report most of the characteristics of the yield curve including the fact that the average yield curve is upward sloping, and that the short end of the yield curve is more volatile and less persistent than its long end. The evidence in Table 2.1 suggests that these two features of the yield curve can be explained by the behavior of risk premiums.
Table 2.1 also shows that highest average risk premiums are in general achieved by EMEs, no matter the maturity. The average $3-$month risk premium ranges between 11.45 basis points (bps) and 34.12 bps in AEs, whereas in EMEs, it ranges between -3.01 bps and 183.74 bps. This reveals some homogeneity among AEs, which contrasts with the situation in EMEs. This observation is the same for longer maturities. Indeed, the average $6-$month (resp. $60-$month) risk premium ranges between 25.63 bps and 59.99 bps (resp. 123.97 bps and 250.35 bps) in AEs, whereas in EMEs, it rangers between 30.75 bps and 161.64 bps (resp. 3.59 bps and 307.70 bps). The observation that average risk premiums in AEs are much less disparate than in EMEs is very intuitive. Indeed, the more similar countries are in terms of their economic fundamentals, political risks, and resilience to external shocks the more homogeneous they will be in terms of risk perceived by investors. This is the case for most developed economies, but not for developing and emerging economies which generally face more idiosyncratic shocks, have higher political risks and greater fiscal strains that lead investors to demand higher compensations.

Table 2.1: Descriptive statistics, forward risk premiums by maturity.[a]

3-Month

	AUS	BEL	CAN	CRO	FRA	HKG	HUN	ITA	MLS	MLT	MEX	NOR	PHI	SAF	SWE	USA
Mean (in bps)	12,31	17,63	30,13	126,74	20,82	31,83	17,40	22,92	12,92	24,31	107,50	11,45	183,74	-3,01	15,12	34,12
Std dev. (in bps)	39,23	38,45	37,86	124,72	36,97	60,33	111,83	38,32	41,22	33,23	193,36	48,54	152,08	57,12	39,89	48,19
Coef. of Var.	3,19	2,18	1,26	0,98	1,78	1,90	6,43	1,67	3,19	1,37	1,80	4,24	0,83	-19,00	2,64	1,41
Minimum (in bps)	-111,40	-41,00	-37,00	-159,73	-37,03	-143,96	-317,00	-41,96	-84,80	-55,00	-149,99	-91,74	-420,08	-169,18	-61,00	-22,00
Maximum (in bps)	192,15	232,00	198,02	435,67	230,02	333,00	420,01	207,00	259,60	186,00	1044,11	287,03	687,05	153,27	249,01	233,07
$\rho(1)$	0,37	0,79	0,63	0,79	0,84	0,76	0,69	0,65	0,71	0,80	0,82	0,73	0,76	0,67	0,82	0,85
$\rho(12)$	-0,23	0,07	0,12	0,13	0,08	0,10	-0,13	-0,01	0,08	0,18	-0,08	0,01	0,19	0,19	-0,10	0,18
$\rho(24)$	-0,09	-0,10	-0,15	-0,17	-0,08	0,08	0,02	0,03	0,14	-0,02	0,50	-0,06	0,17	0,03	-0,27	-0,21

6-Month

	BEL	CAN	CRO	FRA	HKG	HUN	ITA	MLS	MLT	MEX	NOR	PHI	USA
Mean (in bps)	59,99	49,60	101,48	30,66	58,46	30,75	40,98	34,93	35,58	136,15	25,63	161,64	37,55
Std dev. (in bps)	87,31	68,64	138,50	69,97	90,13	177,61	111,19	58,16	55,04	228,40	90,03	188,54	78,41
Coef. of Var.	1,46	1,38	1,36	2,28	1,54	5,78	2,71	1,66	1,55	1,68	3,51	1,17	2,09
Minimum (in bps)	-79,99	-77,98	-285,69	-90,07	-175,85	-588,00	-404,80	-48,69	-79,00	-212,00	-125,95	-385,02	-88,00
Maximum (in bps)	414,77	264,06	503,58	337,87	347,03	513,74	570,90	345,60	269,02	999,03	388,01	842,21	257,00
$\rho(1)$	0,90	0,85	0,86	0,91	0,89	0,84	0,70	0,82	0,84	0,88	0,88	0,84	0,94
$\rho(12)$	0,03	0,19	-0,04	0,07	0,25	-0,26	-0,02	0,06	0,12	-0,05	-0,01	0,13	0,33
$\rho(24)$	-0,17	-0,22	-0,19	-0,19	-0,05	0,00	-0,10	0,23	-0,19	0,58	-0,14	-0,04	-0,30

5-year

	AUS	BEL	CAN	FRA	HKG	HUN	ITA	MLS	MLT	NZL	NOR	SIN	SWE	SWI	UK	USA
Mean (in bps)	123,97	226,12	236,70	241,40	307,70	3,59	175,15	156,84	214,35	129,74	217,49	231,53	222,41	208,97	172,40	250,35
Std dev. (in bps)	129,18	123,71	57,45	110,38	106,66	246,27	132,63	102,80	103,50	107,43	82,03	75,99	82,67	82,33	154,46	106,35
Coef. of Var.	1,04	0,55	0,24	0,46	0,35	68,51	0,76	0,66	0,48	0,83	0,38	0,33	0,37	0,39	0,90	0,42
Minimum (in bps)	-145,91	-46,99	131,02	-9,24	9,02	-597,23	-267,27	-30,01	11,18	-81,85	-16,73	-15,68	31,46	17,23	-75,87	42,06
Max (in bps)	418,06	497,50	372,54	458,52	565,41	743,07	471,04	436,28	482,30	364,01	377,30	354,41	394,49	360,29	454,05	450,01
$\rho(1)$	0,96	0,97	0,85	0,97	0,94	0,85	0,97	0,94	0,98	0,96	0,94	0,87	0,95	0,95	0,98	0,95
$\rho(12)$	0,44	0,49	0,02	0,61	0,24	0,49	0,39	0,31	0,50	0,43	0,31	0,02	-0,05	0,45	0,79	0,55
$\rho(24)$	0,34	0,02	-0,05	0,24	-0,31	0,06	-0,38	0,04	-0,12	0,50	0,08	0,00	-0,04	0,26	0,63	0,26

[a]Countries acronyms are as follows: Australia (AUS), Belgium (BEL), Canada (CAN), Croatia (CRO), France (FRA), Hong Kong (HKG), Hungary (HUN), Italy (ITA), Malaysia (MLS), Malta (MLT), Mexico (MEX), New Zealand (NZL), Norway (NOR), The Philippines (PHI), Singapore (SIN), South Africa (SAF), Sweden (SWE), Switzerland (SWI), United Kingdom (UK), United States (USA). 3-month and 6-month forward risk premiums for Croatia are computed for the period $2003M01-2009M06$. 3-month forward risk premiums for Italy are computed for the period $1999M01-2009M06$ for all others. For all other countries the time span is $1999M01-2012M06$ for 3- and 6-month forward risk premiums, and $1999M01-2010M04$ for 60-month forward risk premiums.

These results also give a preliminary insight about the gap between EMEs and AEs, as far as risks are concerned. One can say from this table that not only average risk in EMEs have been very different from that of AEs, but even within EMEs (that share common economic characteristics) there is a lot of disparity. The following section seeks to discover if, in spite of this observation, the gap between the two types of economies has narrowed over time.

2.4 Convergence of Premiums

In this section, empirical evidence on risk premium convergence between EMEs and AEs are presented. Three complementary[6] statistical and econometric exercises are performed to analyze the convergence of risk premiums between the countries of interest during the period $1999M01 - 2012M06$. First, raw correlations of premiums on bonds of the same maturity are analyzed. Then, principal component analysis is performed on risk premiums of the same maturity for all countries in order to investigate whether there is a main latent factor driving risk premiums. Finally a deeper analysis is performed, that consists in constructing maximum differentials for each maturity (an artificial series which, for each period, reports the highest risk premium spread between EMEs and AEs) and showing results for each artificial series regressed on a time trend and a few other control variables; the significance and the sign of the coefficient multiplying the trend in these regressions determine whether there is convergence or divergence between countries in terms of risk.

2.4.1 Correlations

The results of the first exercise are reported in Table 2.2. Not all countries have data available to compute risk premiums of a given maturity. So, the number of countries differs from one panel of the table to another. Boldface entries show the correlations between Advanced Economies and emerging countries' risk premiums. The top panel of the table presents the correlations for the 3-month forward risk premiums. The difference between the correlations among Advanced Economies and the correlations between advanced and emerging countries is striking. Indeed, the correlations among Advanced Economies are positive and non negligible; the lowest value is 0.32 (between Norway and the U.S.). Not surprisingly, the largest values are reached by country pairs that are in the EMU (Belgium, France, and Italy).

In contrast, the correlations between advanced and emerging countries are low and mostly negative. The highest positive correlations are obtained by Hong Kong (0.55 with the U.S. and 0.42 with Canada). These highest correlations are closer to those found among Advanced Economies because the Hong Kong dollar is pegged to the U.S. dollar, and

[6]The complementarity of these exercises relies on the fact that some of them are static, and others dynamic.

Table 2.2: Correlations of forward risk premiums across countries and by maturity.[b]

3-Month

	BEL	**CAN**	**FRA**	**ITA**	**MLT**	**NOR**	**SWE**	**USA**
BEL	1,00							
CAN	0,48	1,00						
FRA	0,92	0,57	1,00					
ITA	0,79	0,53	0,83	1,00				
MLT	0,59	0,33	0,61	0,60	1,00			
NOR	0,50	0,34	0,56	0,55	0,38	1,00		
SWE	0,61	0,46	0,70	0,69	0,40	0,52	1,00	
USA	0,46	0,77	0,50	0,42	0,35	0,32	0,36	1,00
CRO	**-0,28**	**-0,28**	**-0,29**	**-0,25**	**-0,38**	**-0,10**	**-0,09**	**-0,48**
HKG	**0,24**	**0,42**	**0,27**	**0,23**	**0,15**	**0,11**	**0,08**	**0,55**
HUN	**-0,06**	**-0,12**	**-0,07**	**0,01**	**0,02**	**-0,16**	**0,01**	**-0,17**
MLS	**0,16**	**0,15**	**0,21**	**0,17**	**0,03**	**-0,18**	**0,24**	**0,02**
MEX	**-0,06**	**0,04**	**-0,01**	**0,03**	**0,13**	**-0,10**	**0,01**	**-0,02**
PHI	**-0,18**	**0,01**	**-0,17**	**-0,16**	**-0,01**	**-0,17**	**-0,06**	**0,14**
SAF	**0,16**	**0,07**	**0,17**	**0,13**	**0,07**	**0,21**	**0,01**	**-0,05**

6-Month

	BEL	**CAN**	**FRA**	**ITA**	**MLT**	**NOR**	**USA**
BEL	1,00						
CAN	0,45	1,00					
FRA	0,87	0,61	1,00				
ITA	0,78	0,28	0,60	1,00			
MLT	0,63	0,25	0,67	0,44	1,00		
NOR	0,55	0,46	0,71	0,35	0,55	1,00	
USA	0,48	0,80	0,59	0,33	0,34	0,36	1,00
CRO	**-0,31**	**-0,38**	**-0,39**	**-0,07**	**-0,24**	**-0,47**	**-0,32**
HKG	**0,24**	**0,68**	**0,33**	**0,15**	**0,17**	**0,23**	**0,78**
HUN	**0,00**	**-0,16**	**-0,02**	**0,05**	**-0,03**	**-0,19**	**-0,19**
MLS	**0,12**	**0,29**	**0,24**	**0,05**	**0,13**	**0,17**	**0,14**
MEX	**-0,08**	**0,08**	**-0,11**	**-0,11**	**0,17**	**-0,10**	**0,03**
PHI	**-0,16**	**-0,06**	**-0,09**	**-0,16**	**0,00**	**-0,21**	**-0,01**

60-Month

	AUS	**BEL**	**CAN**	**FRA**	**ITA**	**MLT**	**NZL**	**NOR**	**SWE**	**SWI**	**UK**	**USA**
AUS	1,00											
BEL	0,72	1,00										
CAN	0,70	0,39	1,00									
FRA	0,85	0,94	0,53	1,00								
ITA	0,25	0,78	0,10	0,60	1,00							
MLT	0,51	0,81	0,14	0,76	0,71	1,00						
NZL	0,94	0,65	0,69	0,81	0,21	0,50	1,00					
NOR	0,83	0,73	0,51	0,78	0,36	0,52	0,73	1,00				
SWE	0,85	0,77	0,72	0,83	0,46	0,54	0,81	0,84	1,00			
SWI	0,89	0,78	0,63	0,90	0,35	0,53	0,86	0,77	0,82	1,00		
UK	0,81	0,47	0,68	0,70	-0,02	0,36	0,88	0,51	0,62	0,78	1,00	
USA	0,74	0,27	0,73	0,52	-0,17	0,04	0,79	0,43	0,53	0,66	0,86	1,00
HKG	**0,06**	**-0,13**	**0,18**	**-0,13**	**-0,18**	**-0,41**	**0,04**	**0,17**	**0,14**	**0,01**	**-0,14**	**0,28**
HUN	**0,44**	**0,72**	**0,06**	**0,70**	**0,62**	**0,75**	**0,42**	**0,44**	**0,38**	**0,55**	**0,34**	**0,06**
MLS	**-0,10**	**0,09**	**-0,15**	**-0,02**	**0,32**	**-0,11**	**-0,16**	**0,07**	**0,08**	**-0,02**	**-0,37**	**-0,12**
SIN	**0,49**	**0,18**	**0,42**	**0,30**	**-0,02**	**-0,12**	**0,48**	**0,29**	**0,34**	**0,46**	**0,40**	**0,65**

[b]Correlations for Croatia are based on the subsample $2003M01 - 2009M06$, and correlations for Italy on the subsample $1999M01 - 2009M06$. Boldface entries are correlations between AEs and EMEs risk premia. Missing countries are those for which some interest rate data are missing and, as a consequence, one cannot compute risk premiums for them.

interest rates in Hong Kong are somewhat driven by their U.S. counterparts; this a well-known behavior of interest rates in fixed exchange rate regimes with open capital accounts. The situation is almost similar with Malaysia. Notwithstanding this evidence, most of the correlations for these two countries remain very low compared to what could be expected for countries with their exchange rate regimes [7]. This provides evidence about

[7]Indeed, one expects high correlations of interest rates of two countries when one is pegged to the

the absence of convergence between AEs' 3-months forward risk premiums and their EMEs counterparts.
The results observed for 3-month forward risk premiums are consistent across maturities. The middle panel of Table 2.2 repeats the analysis using 6-month forward risk premiums. The correlations among advanced economies are slightly higher than in the first panel but they remain mostly low and negative between advanced and emerging economies. Hong Kong and Malaysia still display positive correlations with advanced economies due to the peg, but these correlations are in general much smaller than the ones among advanced economies. Finally, the bottom panel of the table shows that the evidence concerning 5-year forward risk premium is mostly the same albeit less strong than for previous risk premiums. The results presented in the middle and the bottom panels of Table 2.2 therefore lead to the same conclusion drawn for the 3-month risk premiums; which is that there is a significant difference in the dynamics of risk premiums in EMEs compared to AEs.

2.4.2 Principal Component Analysis

The second exercise conducted in order to assess the convergence of risk premiums is Principal Component Analysis (PCA) whose aim is to summarize the information included in a large set of variables in a few number of factors. The main objective of this exercise is to assess whether there exists a common latent factor (the first principal component) that drives risk premiums of both advanced and emerging countries. Ehrmann et al. (2011) also uses PCA to assess the convergence of yield curves in the Euro Area.
Let X denote a $T \times n$ matrix with rows corresponding to months and columns corresponding to forward risk premiums of a given maturity for n economies. The PCA allows X to be written as

$$X = F\Lambda + \eta \tag{2.10}$$

where F is a $T \times k$ matrix of unobserved factors (with $k < n$), Λ is a $k \times n$ matrix of factor loadings, and η is a $T \times n$ matrix of white noise error terms. Had risk premia of a given maturity converged, one would expect the first principal component to explain most of the total variation included in the data; in Ehrmann et al. (2011)'s terms, there would exist a $T \times 1$ vector F and constants $\lambda_i,\ i = 1, \cdots, n$ such that the matrix X is described by $F \times [\lambda_1, \cdots, \lambda_n]$ up to an error term.
For each maturity, the percentage of total variation explained by each of the first three principal components is computed and the results are reported in Table 2.3. As already noted, had risk premiums converged between advanced and emerging countries, one would notice that the first principal component explains most of the variation; but specifying a

other. For instance, Ehrmann et al. (2011) study convergence between Germany and the other Euro Area countries by looking at correlations of bond yields during the pre- and the post-EMU periods, they find a minimum correlation of 97% in the pre-EMU period when most countries' currencies were almost pegged to the German currency.

Table 2.3: Share of total variation explained by the first three principal components in a Principal Component Analysis of yields across countries and by maturity

	3-Month	**6-Month**	**60-Month**
Contribution of			
First PC	30,05	32,85	54,95
Second PC	20,38	21,76	22,28
Third PC	13,49	15,38	10,97

threshold above which one could conclude that risks have converged is not readily obvious. In their study on convergence of yields in the Euro Area, Ehrmann et al. (2011) consider that a total variation of 97% explained by the first principal component was not enough to conclude that yields were driven by a single factor in the pre-EMU period. Our analysis shows that for all maturities, the first three principal components explain less that 90% of total variation included in risk premiums (see Table 2.3); and besides, less than 55% is explained by the first principal component.

The second and third columns of the table show that the first principal components of 3- and 6-month risk premia explain only 30.05% and 32.85% respectively (i.e. barely one third) of the total variation. The performance of the first principal component of 5-year risk premiums is much better (54.95%) due to the fact that, as mentioned in correlation analyses, the four EMEs for which data are available have (or have had) their currencies pegged on those of AEs. Nevertheless, this evidence does not allow to conclude that 5-year risks have converged since almost 45% of their variation remains unexplained by the first factor. The analyses in this section clearly indicate that there is there are several factors that drive risk premia in EMEs and AEs, i.e. we cannot conclude that there has been convergence.

Further examination of the results of this exercise reveals that in addition to the fact that the first principal components do not explain much of the variation in the data, the time series associated with them are not decreasing over time (see Figure 2.2). Indeed, for the 3- and 6-month risk premiums, they remain stable over time. However, for the 5-year risk premium, the first principal component has been increasing except for the two year period between 2000M06 and 2002M06. This additional evidence strengthens the previous statement that there has not been convergence between EMEs and AEs.

2.4.3 Maximum differentials and Regression Results

In the third exercise, for each maturity, I construct maximum differentials of risk premiums. This is an artificial time series aiming at capturing the largest spread in risk premiums between advanced and Emerging Market economies. For a given maturity k, the maximum differential between emerging and Advanced Economies at period t is given by

$$D_t(k) = \max\{fp_t^i(k),\ i \in EMEs\} - \min\{fp_t^j(k),\ j \in AEs\} \tag{2.11}$$

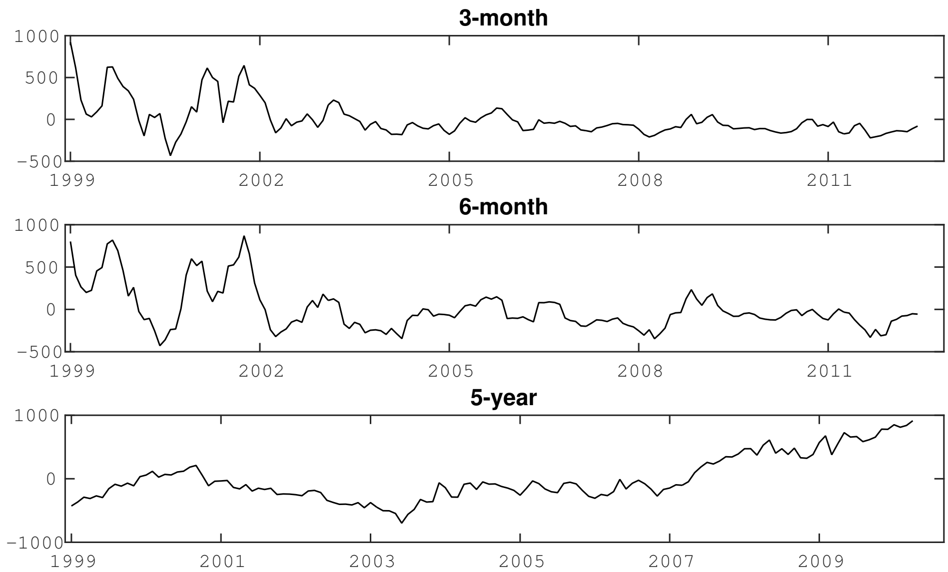

Figure 2.2: Time series of first principal components

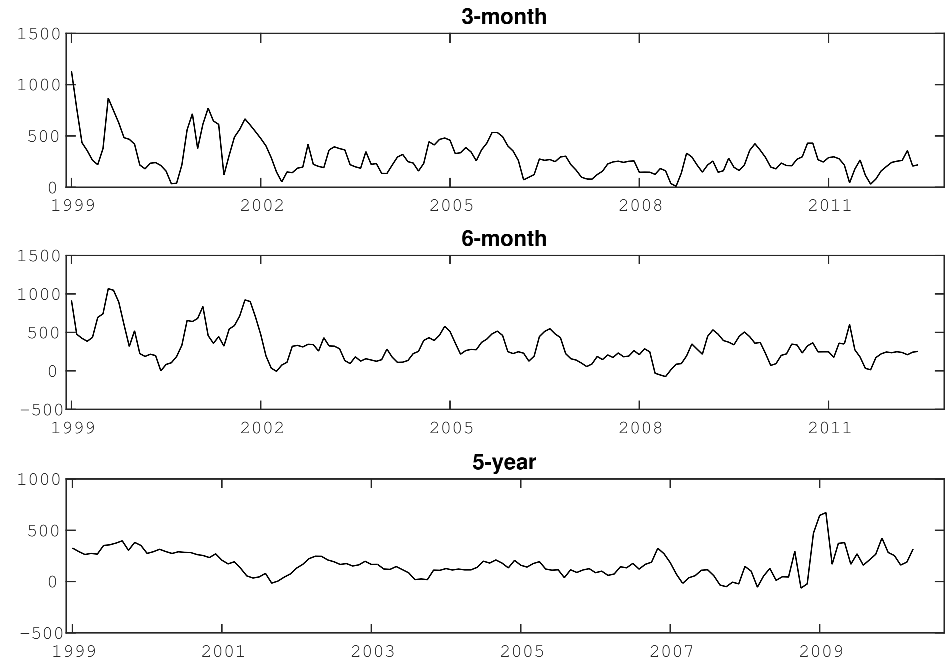

Figure 2.3: Time series of maximum differentials by maturity (in basis points)

To understand the intuition behind maximum differentials, consider two artificial economies, E and A, such that the period t risk premium in E is the maximum of emerging countries' risk premiums, and the period t risk premium in A is the minimum of Advanced Economies' risk premiums. Country E's risk of maturity k represents the worst risk performance an emerging country could have had given observed risk premia; and country

A's risk of maturity k represents the best risk performance an Advanced Economy could have had given observed risk premiums. The development of government bond markets in the direction of greater unification of non-credit risk premiums between EMEs and AEs should lead to a downward trend of the differentials between the risk premiums of countries E and A[8]. This differential is given by (2.11) and plotted in Figure 2.3.
Figure 2.3 shows that for maturities of 3 and 6 months, maximum differentials between EMEs and AEs were high and relatively stable since 2002. Although they were slightly higher and more volatile before 2002, there is visually no downward trend. Concerning the maximum differentials of risk premiums of 5-year maturity, it is apparent that this time series is characterized by three subperiods. The first subperiod is before 2000 where differentials were high (about 475 bps) and stable. The second begins in 2000 when the series decreased gradually and stabilized around 200 bps, till 2008. The last subperiod is after 2008, when the series slightly decreased and began to be very volatile.
An important observation that derives from the analysis of Figure 2.3 is that the small decreases observed in maximum differentials coincide with periods of loose monetary conditions in Advanced Economies (early and late 2000s). The decreases can therefore be associated with lower liquidity risk differentials between EMEs and AEs since in such periods, funds flow towards Emerging Markets as investors seek higher yields.
The latter observation leads to a supplemental exercise that consists in regressing each maximum differential on a time trend, an indicator of global liquidity condition (the Federal Funds Rate - Fedfunds), and an indicator of investors risk aversion (the Chicago Board Options Exchange's volatility index - VIX). The intuition is that after controlling for global factors, if Emerging Markets have been converging, the coefficient multiplying the time trend should be significantly negative. The results of this additional exercise are presented in Table 2.4; the coefficients are estimated using least squares with structural breaks. These are tested using Bai tests of breaks in all recursively determined partitions. As Table 2.4 shows, the coefficient multiplying the time trend is indeed negative in all regressions, but is not significantly different from zero for any of them. This leads to the conclusion that risk premiums did not converge. The regressions show that there have been structural breaks in the time series of 3- and 6-months risk premiums, consistently with the visual impression. Surprisingly, the structural break tests fail to find structural changes in 5-year risk premium series, even after different types of tests are considered.
To summarize this section, Emerging Market countries' credit risk have considerably decreased during the last decades (see Baldacci (2007), González-Rozada & Levy-Yayeti (2008), IMF (2003, 2004), among others), but these countries still lag far behind when the other sources of risk that are important to investors are considered. The three complementary statistical and econometric exercises constitute a robustness check for this finding. The results also do not significantly change during the last episode of financial and economic crisis as shown by some additional robustness checks.

[8]Of course, this makes sense only if risk premiums in Advanced Economies are in general lower than those of Advanced Economies, which is the case in the data.

2.4.4 Convergence Analyses with an Alternative Measure of Risk Premium

In this Section, the convergence analyses are repeated with a measure of risk premium where the expected rates are constructed using averages of current and past realized rates. For instance at time t, the expected rate of a 3-month (resp. 6-month) treasury bill in three (resp. six) months is proxied by the average rate of that asset in the current and the last two (resp. five) months. I focus only on 3- and 6-month maturities in order to have reasonable sample lengths for the different analyses because long time series are not available for some emerging market economies.

The results of this exercise can be found in Appendix B.2. Starting with the correlation exercise, the panels of Table B.6 present similar distinctions as the first two panels of Table 2.2, regarding correlations among Advanced Economies, and between Advanced and Emerging Market Economies. Indeed, the correlation among advanced economies are most of the time very high and always positive, whereas between Advanced Economies and Emerging Markets, they are often smaller and negative. As for the baseline analyses, there are some exceptions for low correlations between EMEs and AEs; these are EMEs that have (or until recently had) a fixed exchange rate regime with an advanced economy (e.g. Malaysia), and whose interest rates were influenced by those of the advanced economies on which they pegged.

The principal component analysis with the alternative measures of forward risk premium shows that the first principal component for each maturity explains a higher share of the total variation than in the baseline (see Table B.7 in Appendix B.2). The first principal component for the exercise with 3-month risk premiums explains 49.46% of the total variation in the data compared to 30.05% for the baseline; these figures are respectively 45.71% and 32.86% for 6-month risk premiums. In spite of the marked improvement in these numbers, the shares explained by the first principal components remain small relative to the share that could be considered as a evidence of convergence (e.g. more that 98% as in Ehrmann et al. (2011)). Besides, as Table B.7 shows, the second and third principal components explain a non-negligible of the total variation in the data, for each maturity; which provides additional evidence that there is more than one factor that drives forward risk premiums in the two categories of economies studied.

Table 2.4: Regression results (Maximum differentials).[c]

	fp(3)		fp(6)			fp(60)
	1999M01-2002M02	2002M03-2012M06	1999M01-2001M12	2002M01-2008M03	2008M04-2012M06	
Constant	133.39	340.46***	936.73	187.21	1116.24	279.72***
	(450.83)	(93.32)	(774.23)	(166.80)	(753.22)	(98.01)
Trend	**-8.29**	**-0.48**	**-6.72**	**-0.12**	**-5.45**	**-0.74**
	(6.31)	**(0.55)**	**(10.56)**	**(2.44)**	**(4.55)**	**(0.87)**
Vix	31.00***	-1.67	11.72	1.78	-1.41	0.52
	(11.64)	(1.77)	(11.40)	(3.89)	(4.14)	(2.51)
Fedfunds	-71.93*	-6.24	-119.86	19.94	-217.66***	19.72
	(39.83)	(17.98)	(79.04)	(50.46)	(75.44)	(18.63)
Obs.	162			162		136
R^2	0.43			0.49		0.25

[c]Numbers in parentheses are robust standard errors.
*p < 10%. **p < 5%. ***p < 1%.

Finally, a closer look at the artificial time series of maximum differentials based on the alternative measures of risk premiums (Figure B.1 in Appendix B.2) shows two apparent phases in the dynamics of the series. A phase of gradual decline from the 1999M01 to 2007M02 for the 3-month maturity and the beginning of 2003M01, and a phase of relative stability afterwards. Table B.8 shows the results of a regression similar to the one done in Section 2.4.3. After controlling for global macroeconomic and financial conditions, it is still impossible to say with certainty that forward risk premiums of short maturities in EMEs have converged to AEs' levels. The regression based on 3-month forward risk premiums on bonds do show significantly negative coefficients. However for the 6-month maturity, although the coefficients are all statistically significant, they are positive when considering some sub-samples meaning that not only did risk premiums not converge, but in fact, the difference even increased between 2002 and 2005, and after 2009.

The evidence based on the alternative measure of expected rates on bonds is not as clear as the baseline; as just seen the maximum differentials show a slightly decreasing trend for the 3-month risk premiums but also periods of increased differentials for 6-month risk premiums. However, the first two exercises performed in this section reinforce the findings of the preceding sections that there has not been convergence. The remainder of the paper investigates the determinants of risk premium differentials for the maturities enumerated so far. It focuses on the role played by global factors as well as differences in the macroeconomic fundamentals and political risks between EMEs and AEs.

2.5 Determinants of Risk Premium Differentials

As shown in the previous section, non-default risk premium differentials between Emerging Market economies and Advanced Economies have been quite large and stable during the period $1999 - 2012$. The present section investigates how countries' macroeconomic fundamentals, the political environment, as well as global economic and financial conditions have influenced these differentials.

2.5.1 Empirical Model Specification, Data and Methodology

The theoretical model presented in Section 2.2.2 provides a framework of analysis for the empirical investigations of the factors that drive forward risk premium differentials between EMEs and AEs. For a given maturity, k, the following panel model is specified:

$$y_{ijt} = \beta_0 + \beta_1 x_{ijt} + \beta_2 z_t + \alpha_{ij} + \epsilon_{ijt} \tag{2.12}$$

In (2.12), the cross-section dimension of the panel represented by the subscript ij refers to a country pair composed of Emerging Market economy $i \in \{1, \cdots, 8\}$ and Advanced Economy $j \in \{1, \cdots, 10\}$. I consider all possible emerging-advanced country pairs and, thus, do not consider any Advanced Economy as a benchmark; this will allow me to later study how regional unobserved factors influence risk premium differentials. y_{ijt}

denotes the difference between the forward risk premium of Emerging Market economy i and that of Advanced Economy j in period t. x_{ijt} denotes a vector of explanatory variables specific to each country pair (i, j); each variable in x_{ijt} is the ratio or difference in values of some macroeconomic indicator, between Emerging Market i and Advanced Economy j. z_t denotes a vector of global factors that affect y_{ijt}. β_0, β_1 and β_2 are the coefficients that capture the effects of the explanatory variables; α_{ij} represents an effect that is specific to country-pair (i, j); and ϵ_{ijt} is a white noise error term assumed to be uncorrelated with explanatory variables and with the pair specific effects. (2.12) therefore models the forward risk premium differential as a function of countries' differences in their fundamentals and of global factors, with possible pair specific effects.

For the benchmark model, the variables included in x_{ijt} are: *(i)* the inflation rate differential, *inflation*, as well as the GDP growth rate differential, *growth*. Inflation and economic growth are perceived by investors as indicators of macro-economic stability. High inflation differentials are expected to increase risk differentials because they reduce the real wealth of investors. In contrast, high growth differentials lead to lower risk on the part of Emerging Markets; *(ii)* the relative level of Emerging and Advanced Economies in terms of political risk, *polrisk*. I use a composite index of political risk, namely the ICRG political risk variable produced by the PRS Group. This index takes into account factors like the political and governmental stability, the investment profile of the country, governance, the presence of conflicts, etc. A higher degree of political risk in a country relative to others is expected to increase the risk differential[9]; *(iii)* the relative level (ratio) of Emerging and Advanced Economies in terms of the liquidity of their financial market, *liquid*. To proxy liquidity, I use the ratio of a country's outstanding government securities to the total of outstanding government securities for the whole sample of countries included in the analyses. More liquid emerging financial markets are a good sign for investors, and are therefore expected to entail less risks; *(iv)* the real effective exchange rate ratio between Emerging and Advanced Economies, *reer*. The expected effect of *reer* on the risk differential is unclear since it can be the result of two conflicting effects. First, a real appreciation of the currency leads to less competitive exports and can trigger economic instability, especially for Emerging Markets. Such a situation is considered as risky by investors who, consequently, increase risk premiums. Second, a real appreciation increases the purchasing power of the home currency abroad, and is therefore perceived by (foreign) investors as advantageous because it increases the cash flow from their investment measured in a foreign currency.

Two variables are included in z_t and are supposed to capture the global financial and economic environment, namely the Fed Funds rate (*fedfunds*) and the Chicago Board Option Exchange volatility index (*vix*). High values of *fedfunds* stem from adverse international liquidity conditions (Baldacci et al. (2008)); increases in this variable are therefore expected to increase risk premium differentials. By including the *vix* indicator in the model,

[9]The political risk measure is designed so that an increase represents an improvement, i.e. a less risky political environment. The index ranges between 1 and 100.

I particularly aim to capture how risk aversion vis-à-vis Advanced Economies' securities affects the risk premium differential between EMEs and AEs. An increase is expected to lower the risk premium differential between EMEs and AEs because investors will increase their demand for Emerging Market securities due to high risk in AEs, which decreases the risk spread.

Data on inflation, GDP growth and real exchange rate data are from the IMF's International Financial Statistics (IFS) database. In the regression using short-term risk premium differentials, I use current inflation and growth rates, while for the regression on the five-year risk premium differential, I use their five-year expected values obtained by averaging the current and past realized rates over the last five years. An increase in the measure of real effective exchange rates used denotes a real appreciation of the corresponding currency. Raw data used to construct liquidity series come from the Bank for International Settlements (BIS). The Fed Funds rate series come from the St-Louis' FRED database and the *vix* data come from Bloomberg.

The monthly series of risk premium differentials are converted to quarterly series by taking end-of-period values. This allows me to avoid the problem of simultaneity that would invalidate our results otherwise. All other variables are also available at the quarterly frequency, but are chosen to be either period-averages or beginning-of-period values. The Breusch-Pagan LM test for random effects versus ordinary least squares rejects the latter in favor of the former, meaning that the model should account for the individual-specific effects. The Hausman test for fixed versus random effects model suggests that the preferred model specification is the latter, meaning the the pair-specific effects α_{ij} should be assumed random. The error term in the model (2.12) is therefore $u_{ijt} = \alpha_{ij} + \epsilon_{ijt}$ where both terms of u_{ijt} are assumed *iid* and uncorrelated with each other.

Table 2.5 shows the descriptive statistics on the time series presented above. Average risk premiums decrease with maturity, while their volatility is an increasing function of maturity. The average political risk ratio between EMEs and AEs is 0.87 with a standard deviation of 0.10. The average liquidity ratio is 1.67, which reflects the fact that EMEs issue more short-term debt as a percentage of their GDP than AEs. However, liquidity ratios are highly volatile (with a coefficient of variation of almost 3). On average, inflation and GDP growth in EMEs more than 1.5 percentage points higher than the values of AEs, but the disparities of these differentials across country-pairs are very high. Finally, real effective exchange rate ratios between EMEs and AEs are closed to one and not very volatile.

A potential problem with the estimation of the present model is that errors can be correlated across country pairs. This can arise in two ways. First, given an advanced economy i, errors associated with country pairs (i, j) and (i, k) are likely to be correlated due to the fact that advanced economy investors' decisions to invest in emerging market country j are not completely independent of their decisions to do so in emerging market country k. Second, the attractiveness of emerging market economy j for investors from different advanced economies are very likely to be the same at a given moment in time, so that

Table 2.5: Descriptive statistics for the panel data.[d]

	Description	Mean	Standard deviation	Min	Max
Risk 3-M	3-month risk premium differentials	43.74	138.28	-435.01	771.12
Risk 6-M	6-month risk premium differentials	39.49	185.50	-665.64	1047.57
Risk 5-Y	5-year risk premium differentials	30.01	201.06	-897.53	536.04
polrisk	Political risk ratio	0 .87	0.10	0.63	1.18
liquid	Liquidity ratio	1.67	4.93	0.00	51.33
inflation	Inflation differential	1.59	3.33	-9.13	18.64
growth	GDP growth differential	1.67	3.28	-10.28	17.62
reer	Real effective exchange rate ratio	1.02	0.20	0.50	2.26
vix	CBOE volatility index	23.11	9.45	11.04	61.18
fedfunds	Fed Funds rate	2.61	2.14	0.07	6.54

[d]Ratios and differentials are between EMEs and AEs..

errors associated with country pairs (i, j) and (l, j) are most probably correlated as well. Tackling this issue is crucial for good statistical inference, I estimate two-way cluster-robust standard errors (see Cameron et al. (2011) and Cameron & Miller (2014*a*)); the groups in the first dimension of clustering are indexed by individual advanced economies and those in the second dimension are indexed by individual emerging market countries[10]. An additional issue is the small number of clusters in each dimension (12 for the first dimension and 8 for the second) which can still lead to over-rejection of the null hypothesis when conducting statistical inference. The literature has not yet proposed a way to tackle this issue, but Cameron et al. (2011), Cameron & Miller (2014*a*) and Cameron & Miller (2014*b*) suggest using a T distribution with degrees of freedom equal to the minimum between the number of clusters (eight degrees of freedom in our case) instead of a normal distribution. I follow this recommendation.

The baseline model given by (2.9) is estimated and the results are presented in the following tables, along with the results of many extensions of the baseline specification. Particularly, structural breaks are considered in order to investigate how the recent phases of financial and economic turmoil have shifted the way investors react to changes in global conditions. The interactions between the global risk appetite factor and country-specific fundamentals are also considered in order to investigate the potential amplifying effects of adverse global conditions in investors' responses to countries fundamentals. This is done through the introduction of slope dummies for periods where *vix* exceeds 25 (i.e. periods of high investors' risk aversion; see Baldacci et al. (2008)). Finally, regional dummies are included into the model in order to capture potential geographical discrimination by

[10]For more insights about the problems associated with within cluster error correlation and with few clusters, see Cameron et al. (2011) and Cameron & Miller (2014*a*). Note that unlike residual terms of regressions based on data on bilateral trade, the residuals of our model do not have a complicated pattern of error correlation. The relations between two countries here are unilateral, from an advanced economy to an emerging market economy. So, there is not need to consider robust inference for dyadic data as described in Cameron & Miller (2014*b*); two-way clustering is the appropriate way to tackle potential correlation between the error terms

investors. As robustness checks, the model is estimated using the alternative measures of risk premium defined in Section 2.4.4; to deal with potential endogeneity issues[11], Two-Stage Least Squares estimators are also used on the different extensions of the model with instruments being lagged explanatory variables.

2.5.2 Estimation Results and Robustness Analysis

Estimation Results

Tables 2.6, 2.7 and 2.8 present the results of the baseline specification of the model as well as four extensions, for the 3-month, 6-month and 5-year maturities respectively. The models are estimated over the time period 1999Q1-2012Q2 for maturities 3-, and 6-month; and 1999Q1-2010Q1 for maturity 5-year. For each maturity, the countries considered are the same as in Table 2.2. Model (1) is based on the baseline specification given by equation (2.12). Extension (2) includes structural breaks on the slope coefficients as well as on the intercepts; these structural break are included by using dummy variables that aim to isolate the effects of the crisis of the early 2000s (*dum2002*), and those of the last economic and financial crisis (*dum2008*). Recall from the regression analysis conducted in Section 2.4.3 that the "maximum differential" artificial time series featured structural breaks; this extension of the baseline model deepens the analysis by investigating the impact of these structural breaks on response of risk premium differentials to our explanatory variables. Extension (3) considers the amplifying effects of high market volatility. Extension (4) includes regional effects; the regional dummies are *euro, europe, nafta* for advanced economies in the European Economic and Monetary Union (EMU), in the European Union (EU) and in the North-American Free Trade Agreement (NAFTA). There are also regional dummies *emeurope* and *emeasia* for EMEs in Europe and Asia. Extension (5) considers all variables that have significant effects in the previous four specifications. Numbers in parentheses are two-way cluster-robust standard errors.

Tables 2.6 and 2.7 show coefficients with theoretically consistent signs for the baseline specification, i.e columns labeled *(1)*, but not all statistically significant. However, table 2.8 shows that inflation is significant with the wrong sign in spite of the fact that 5-year inflation expectations are used to in the model with 5-year risk premiums. The sign of the response of risk differences to inflation is difficult to explain.

Liquidity of the local financial markets and economic growth are the only coefficients that are statistically significant for all maturities in the baseline specification, i.e. model (1). Liquidity matters, as an improvement in the relative liquidity of emerging bond markets reduces risk premiums of all maturities. The size of the effect on 5-year risk premiums is at least three times that of shorter maturities. Indeed a hundred basis points (bps) increase in the liquidity ratio lowers risk premiums differentials respectively by 2.38 bps and 1.48 bps for 3- and 6-month maturity bonds, and by 7.12 bps for 5-year bonds. Similarly, higher growth differentials lead to lower risk differentials for all maturities. A hundred

[11] When talking about endogeneity, I am mostly referring to simultaneity.

Table 2.6: Determinants of risk premium differentials: 3-month maturity.

	(1)	(2)	(3)	(4)	(5)
polrisk	-77.94	-148.59**	-123.72	-315.57**	-108.10
	(82.69)	(50.10)	(108.84)	(102.29)	(64.34)
polrisk*dum2002		-43.87			
		(196.30)			
polrisk*dum2008					
polrisk*vix25			-3.33		
			(37.35)		
liquid	-2.38**	-1.83**	-1.83**	-1.62	-2.30**
	(0.71)	(0.67)	(0.73)	(0.92)	(0.71)
liquid*dum2002		0.76			
		(0.85)			
liquid*dum2008					
liquid*vix25			0.57		
			(0.47)		
inflation	8.46	-1.75	6.89	9.52**	1.15
	(5.01)	(2.31)	(5.65)	(4.90)	(2.27)
inflation*dum2002		17.49***			9.72**
		(2.97)			(3.66)
inflation*dum2008					
inflation*vix25			3.71		
			(3.27)		
growth	-7.30**	-7.22**	-6.69	-8.31**	-6.84***
	(2.52)	(2.03)	(3.74)	(2.93)	(1.78)
growth*dum2002		2.97			
		(4.28)			
growth*dum2008					
growth*vix25			-1.05		
			(4.21)		
reer	77.02	-83.38	46.19	65.53	
	(77.35)	(100.04)	(87.43)	(80.66)	
reer*dum2002		225.89**			31.12
		(84.46)			(16.68)
reer*dum2008					
reer*vix25			59.64		
			(37.08)		
vix	-0.81	5.18	-3.36***	51.27	-1.04
	(0.57)	(3.41)	(0.82)	(52.21)	(0.57)
vix*dum2002		-0.39			
		(5.28)			
vix*dum2008					
fedfunds	3.18	-4.20	1.02	76.25	
	(4.55)	(8.52)	(4.39)	(43.71)	
fedfunds*dum2002		-320.12			
		(197.34)			
fedfunds*dum2008					
euro				38.06**	14.00**
				(13.70)	(6.53)
europe				-11.17	
				(11.48)	
nafta				-10.92**	-17.36***
				(4.02)	(3.56)
emeurope				51.27	
				(52.21)	
emeasia				76.25	
				(43.71)	
dum2002		-320.12			
		(197.34)			
dum2008					
Constant	43.94	287.51**	157.44	212.39	159.04**
	(150.63)	(147.41)	(187.71)	(169.39)	(74.31)
N	3174	3174	3174	3174	3174
Overall R-Square	0.06	0.13	0.10	0.14	0.10
Wald Chi-Square	1046.73***	4234.99***	959.73***	1071.07***	158.43***

The model is estimated over the time period 1999Q1-2012Q2. Model (1) is the baseline specification given by equation (2.12); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

Table 2.7: Determinants of risk premium differentials: 6-month maturity.

	(1)	(2)	(3)	(4)	(5)
polrisk	-268.72**	-405.52*	-296.73**	-258.44**	-396.78**
	(95.40)	(166.35)	(85.59)	(68.86)	(99.07)
polrisk*dum2002		351.03**			378.65**
		(132.94)			(142.53)
polrisk*dum2008		-77.81			
		(139.51)			
polrisk*vix25			20.16		
			(79.18)		
liquid	-1.48*	-1.92*	-1.60*	-1.40	-1.62**
	(0.61)	(0.92)	(0.67)	(0.98)	(0.59)
liquid*dum2002		2.24			
		(2.04)			
liquid*dum2008		0.86			
		(0.75)			
liquid*vix25			1.34*		0.93
			(0.55)		(0.48)
inflation	15.10*	3.84	16.12	14.52	3.21
	(6.69)	(6.68)	(9.43)	(7.74)	(5.14)
inflation*dum2002		17.37*			18.79**
		(7.59)			(5.35)
inflation*dum2008		0.97			
		(6.77)			
inflation*vix25			-3.85		
			(7.19)		
growth	-9.51*	-0.80	-2.81	-12.42**	-8.55
	(4.61)	(6.12)	(5.73)	(4.78)	(4.56)
growth*dum2002		-2.46			
		(9.56)			
growth*dum2008		-12.31			
		(6.75)			
growth*vix25			-15.49		
			(7.83)		
reer	83.09	-48.30	57.71	55.67	
	(66.34)	(56.01)	(91.02)	(74.21)	
reer*dum2002		259.09**			201.87**
		(87.11)			(77.61)
reer*dum2008		-46.23			
		(285.29)			
reer*vix25			46.63		
			(78.33)		
vix	-1.27	5.41	-2.83	-20.15	
	(1.05)	(3.13)	(1.41)	(41.54)	
vix*dum2002		7.04			
		(7.05)			
vix*dum2008		3.33			
		(7.78)			
fedfunds	12.67**	-100.67**	11.39*	39.52	14.33**
	(4.58)	(26.78)	(4.99)	(31.41)	(5.00)
fedfunds*dum2002		-692.71***			-3.80
		(165.82)			(7.62)
fedfunds*dum2008		19.29			
		(284.02)			
euro				8.07	
				(8.88)	
europe				-	
				-	
nafta				-11.37	
				(10.67)	
emeurope				-20.15	
				(41.54)	
emeasia				39.52	
				(31.41)	
dum2002		-692.71***			-555.73**
		(165.82)			(140.29)
dum2008		19.29			
		(284.02)			
Constant	183.69	505.84*	245.02	199.76	360.19***
	(173.04)	(227.16)	(187.05)	(179.94)	(88.43)
N	2156	2156	2156	2156	2156
Overall R-Square	0.13	0.13	0.10	0.14	0.05
Wald Chi-Square	151.32***	245.59***	334.18***	5406.92***	116.40***

The model is estimated over the time period 1999Q1-2012Q2. Model (1) is the baseline specification given by equation (2.12); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

bps increase in the growth rate differential leads to a 7.30 bps decrease in 3-month risk premiums, a 9.51 bps decrease in 6-month risk premiums and a 24.38 bps decrease in 5-year risk premiums.

An increase in the difference in terms of inflation between EMEs and AEs leads to higher 6-month risk differentials, with a one percentage point increase in this variable leading to a 0.15 percentage point increase. The results of model (1) also indicate that higher political risk differentials significantly decrease non-credit risk premium differentials between Emerging Markets and AEs for 6-month and 5-year risk premiums; this increase in political risk differentials could be seen as a relative improvement (or less deterioration) in political stability in Emerging Markets compared to AEs. The impact of political risk on risk premium differentials is considerably higher on 5-year than on 6-month risk premiums. Relative real exchange rate developments seem not to play an important role in investors' asset pricing strategies, except occasionally during crisis. Finally, global factors matter to some extent. The results of the baseline specification show that although higher investors' risk aversion vis-à-vis Advanced Economies' financial markets (as captured in the model by *vix*) does not matter much. Adverse global liquidity conditions (as measured by higher values of the Fed Funds rate) lead investors to increase 6-month and 5-year risk premium differentials. The risk differentials of 6- and 5-year maturities increase respectively by 0.13 percentage point and 0.18 percentage point following a one percentage point increase in the Fed Funds rate.

Let's now see how the results presented above change when some relevant factors such as crisis, high market volatilities and regional factors are accounted for. An important observation here is that the explanatory powers of most of the extended models significantly improve relative to the baseline.

Extension (2), which includes slope and intercept dummies for periods of economic crisis, shows that inflation particularly mattered during the early 2000s' economic turmoil; long-term growth perspectives were also highly valued by investors during the last financial crisis (see results of the model using 5-year premiums). Real exchange rate movements seem to have mattered as well for investors interested in short-term assets during the crisis of the early 2000's. Indeed, concerns about economic stability have led foreign investors to hugely increase 3- and 6-month risk premiums in EMEs in response to a real appreciation during that period whereas in normal circumstances, they seem to not significantly price exchange rate movements. The results show that investors this was not the case during the last financial and economic crisis. For long-term assets, the results show that concerns about macroeconomic stability, and generated by an exchange rate appreciation, lead to an increase in emerging market risk premiums. Finally, it is worth mentioning that the slope dummies for *fedfunds* are significantly negative; which is consistent with theory since easy global monetary conditions are usually observed in periods of global economic downturn.

The impact of political risk on short-term forward risk premiums differentials is amplified by the introduction of structural breaks in the model (extension 2). Indeed, when the

Table 2.8: Determinants of risk premium differentials: 5-year maturity.

	(1)	(2)	(3)	(4)	(5)
polrisk	-648.98**	-144.08	-610.32*	-223.66	-267.23
	(187.45)	(143.31)	(219.81)	(179.22)	(120.36)
polrisk*dum2002		-320.56			
		(172.65)			
polrisk*dum2008		840.30*			-50.58
		(316.89)			(90.91)
polrisk*vix25			41.39		
			(68.41)		
liquid	-7.12**	-0.15	-5.33	-5.22*	-1.75
	(1.98)	(1.67)	(3.17)	(1.96)	(2.54)
liquid*dum2002		-9.68*			-7.76
		(3.86)			(4.89)
liquid*dum2008		-0.42			
		(2.47)			
liquid*vix25			-5.07*		-0.56
			(2.14)		(2.93)
inflation	-15.16***	-48.29***	-18.70***	-10.82**	-45.95**
	(1.41)	(2.59)	(2.63)	(2.07)	(11.21)
inflation*dum2002		32.80***			26.61*
		(4.14)			(8.99)
inflation*dum2008		68.21**			65.14**
		(12.93)			(17.37)
inflation*vix25			4.76		
			(3.17)		
growth	-24.38***	26.37*	-16.34***	-33.32**	13.97**
	(3.93)	(10.90)	(2.72)	(8.46)	(3.34)
growth*dum2002		-31.30			
		(16.43)			
growth*dum2008		-67.97*			-16.25*
		(26.30)			(6.33)
growth*vix25			-20.90*		-11.28**
			(7.99)		(2.71)
reer	66.96	72.65	103.71*	65.53	120.30*
	(35.65)	(84.70)	(38.54)	(49.40)	(45.62)
reer*dum2002		34.95			
		(91.59)			
reer*dum2008		293.73			
		(228.35)			
reer*vix25			-55.98		
			(47.19)		
vix	0.95	-2.18	3.40	-63.78	
	(3.06)	(1.50)	(4.41)	(47.12)	
vix*dum2002		-22.08			
		(13.45)			
vix*dum2008		42.47			
		(21.79)			
fedfunds	18.21**	80.61**	19.94**	114.96	7.95
	(3.33)	(24.63)	(3.55)	(73.50)	(19.41)
fedfunds*dum2002		596.28**			1.86
		(175.54)			(22.57)
fedfunds*dum2008		-1056.82*			29.88
		(443.29)			(26.91)
euro				22.23	
				(25.22)	
europe				-5.91	
				(22.00)	
nafta				-67.96***	-66.84***
				(7.12)	(8.43)
emeurope				-63.78	
				(47.12)	
emeasia				114.96	
				(73.50)	
dum2002		596.28**			60.39
		(175.54)			(60.26)
dum2008		-1056.82*			
		(443.29)			
Constant	493.78**	-21.23	361.98*	51.18	40.17
	(118.00)	(120.10)	(141.05)	(114.95)	(72.58)
N	2160	2160	2160	2160	2160
Overall R-Square	0.30	0.60	0.34	0.43	0.54
Wald Chi-Square	12065.30***	2762.92***	361.80***	361.48***	3834.74***

The model is estimated over the time period 1999Q1-2010Q1. Model (1) is the baseline specification given by equation (2.12); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

baseline specification is extended by including structural breaks, the elasticity associated with *polrisk* goes from a non-significant value of -81.88 bps to a statistically significant value of -148.59 bps for 3-month risk premiums; and from -268.72 bps to -405.52 bps for 6-month risk premiums[12]. Some puzzling results in this specification, however, are the effects of global liquidity conditions (measured by the Fed Fund Rate) and intercept dummies *dum2002* and *dum2008* on risk premiums. These variables appear to have decreasing effect on risk differentials whereas they are expected to have an increasing effect.

When the model is extended to account for the importance of explanatory variables in periods of high market volatility characterized by high investors' risk aversion in AEs (extension 3), the results show that they are less concerned about liquidity, growth and inflation in these periods except, to some extent, when long-term investments are concerned. This evidence is in line with what is expected from investors in a "hunt for yields" that pushes them to take more risk than in normal times.

The next extension (extension 4) investigates region-specific effects in investors' pricing of risk, by considering different groups of AEs as well as different groups of EMEs. The explanatory power of the model improves in this extension relative to the baseline specification. Before commenting on the regional specific effects, note that the inclusion of regional dummies leads *polrisk* to have a significant effect on the 3-month risk differential whereas its significant effect on the 5-year risk differential vanishes. According to the estimates, risk premium differentials on 3-month bonds are 38.06 bps higher with countries from the EMU and 10.92 bps lower for countries from the NAFTA. 5-year risk differentials are also lower by 67.96 bps with NAFTA countries compared to the others. But investors from all regions seem not to discriminate in terms of the risks charged on assets from different EMEs. So to summarize, the models reveal that all things being equal, risk differential will be the same for (South Africa, Canada) and (Singapore, Canada) country pairs, but the risk differential for the (South Africa, Canada) country pair are generally lower than that of the (South Africa, France) country pair.

The last extension (5) considers all the explanatory variables that are significant in the previous specifications. For the effects of real exchange rate movements in this specification, the conclusions remain unchanged for all maturities. Now consider 3-month risk premium differentials. Liquidity, economic growth and regional dummies continue to have significant effects that are comparable to the ones presented previously. Inflation seems to matter only in periods of economic turmoil and the effect of political risk is no longer significant. Turning to the results based on 6-month risk premiums, note that the effect of political risk is considerably amplified and growth does not have a significant effect any longer. For the 5-year risk premiums, the "sign" problems mentioned above are exacerbated and most of the effects are difficult to explain. However, one can say that inflation and growth were important factors taken into account by investors when pricing risk on

[12]It is worth noting that the crisis of the early 2000s dampened this effect of improved political risk differentials by 351.03 bps so that the resulting decrease in differentials during this period was only 54.49 bps following a hundred percentage point increase in *polrisk*.

long-term assets during periods of economics turmoil. Also, risks differentials still appear to be lower between EMEs and AEs from North America.

Robustness Analysis

Although the different specifications of the model presented above represent a robustness exercise on their own, a number of supplemental analyses have been done in order to ensure the robustness of the results just presented. Two additional exercises are conducted: the first exercise estimates the regression on 3-month and 6-month risk premiums using the alternative measures of policy uncertainty presented in Section 2.4.4. In the second exercise, I perform the estimation using a Two-Stage Least Square estimator in order to tackles potential endogeneity issues.

Tables B.9 and B.10 show the results of the first robustness exercise. As in the main results presented above, the regression on the alternative measure of 3-month forward risk premium differentials show that the liquidity of the local financial markets and macroeconomic stability matter most for investors. An increase in the liquidity ratio is followed by a significant decrease in the risk premium differentials; similarly for an increase in the growth rate differential. A noticeable difference with the previous results is the fact that inflation appears to be significant in specification (1); a one percentage point increase in the inflation rate difference between EMEs and AEs leads to increases in risk premium differentials that range between 26 and 44 bps depending on the specification. Another fact that is consistent with the main results is related to the responses of risk differences to movement in exchange rates; again, this response is present only during crisis. All the above is also observed for the regression on 6-month risk premiums, with the additional fact that political risk difference also plays a non-negligible role.

Estimating the model using instrumental variable techniques yields similar results to the main findings. I use one lag of country-pair-specific variables as instruments in order to deal with the fact that these variables may be endogenous. I use the Two-Stage Least Square estimator to estimate the model again on risk differentials of 3-month, 6-month and 5-year maturities. The results of this exercise are presented in Tables B.11, B.12 and B.13. Again, differences in political risk, liquidity, and growth are shown to play important roles in driving the large risk spreads observed between EMEs and AEs. The wrong sign of the effect of inflation in the regression with 5-year rates appears here as well. The impact of global factors do not change considerably compared to what appeared in the main results.

2.6 Conclusion and Policy Implications

This paper investigates the convergence between 8 Emerging Market economies and 12 Advanced Economies as far as risk premiums on factors other than default are concerned. It considers the forward risk premium on government securities and shows through a simple two-period portfolio choice model that the forward risk premium differential between two countries accounts for all potential sources of risk priced by investors except for credit risk. The paper then shows that "non-credit" risk premiums in Emerging Market Economies have not converged to the levels seen in Advanced Economies during the period 1999-2012, a result that contrasts with the very well developed literature on Emerging Markets' spread compression. Using a panel specification based on the theoretical model, the paper also investigates the determinants of non-credit risk premium differentials between Emerging Markets and Advanced Economies. The results show that differences in liquidity and growth across countries are the main drivers of risk premium differentials. Political risk plays a non-negligible role in shifting the risk differentials between these two types of economy because investors highly value political stability, especially in the long-run. Macroeconomic stability is also well rewarded by investors, especially in periods of high market volatility and during crisis. In periods of turmoil such as financial or economic crisis, investors reduce risk premiums charged on bonds from emerging market countries that take steps to ensure exchange rate stability. Such situations are usually shown to cause exchange rate appreciations in EMEs, with in turn translate into increased risk premiums as investors are worried by the macroeconomic stability implications of currency appreciations for the countries. Changes in global factors such as increased investors' risk aversion and easy global liquidity conditions are shown to be beneficial to Emerging Markets.

A number of policy implications for Emerging Markets' policymakers can be derived from the results presented in Sections 3.3 and 2.5. First of all, the results clearly point to the fact that risk premiums in EMEs have not converged to AEs' levels during the period that the study covers, in spite of the fact that it is widely acknowledged that Emerging Markets have been considerably improving their fundamentals in the last few decades. This means that ensuring this convergence remains a great challenge for policymakers in Emerging Markets. Taking measures in this regard will particularly be very beneficial in the present context of normalization of monetary policy in the U.S. where interest rates are rising and global financial conditions are tightening. We have already been witnessing capitals flowing out of EMEs, meaning that risk premiums are rising again. Taking measures to reduce risk on factors other than default will provide investors with additional incentives to maintain their investments in Emerging Markets. These measures could be in the direction of greater financial markets' deepening through improved liquidity as we have seen that it plays an important role in increasing risk differentials.

The investigations on the determinants of non-credit risk premium differentials show that in spite of the role played by global liquidity conditions and investors' risk appetite, local fundamentals (and in particular political risk, liquidity and growth) are very important for reducing the investors' perception of risk, and therefore governments' borrowing costs. This suggests that Emerging Markets' policy authorities should not focus solely on ensuring the solvency of their debt. Apart from improving their credibility in order to reduce default risks, they should further strengthen local fundamentals, especially regarding political risks, economic stability, and exchange rate developments. They should take measures to increase the depth and the liquidity of local capital markets with the aim of reducing other sources of risks, beyond default risk. Moreover, the results show that in periods of financial and economic turmoil, investors pay particular attention to developments in real exchange rates. So, macroeconomic stability is their one of the main concern in such environments, and policy much be implemented accordingly.

Chapter 3

News Shocks, Business Cycles, and the Inflation Puzzle

3.1 Introduction

A long-standing and fundamental question in macroeconomics is: what causes business-cycle fluctuations? Following the seminal work of Beaudry & Portier (2006), interest has been rekindled in Pigou (1927)'s theory of business cycles, according to which changes and revisions in expectations about future fundamentals can give rise to boom-bust cycles. A number of empirical studies — based on vector autoregressions (VARs) — have therefore attempted to gauge the importance of news shocks about future productivity in generating the type of positive comovement of macroeconomic aggregates observed in the data and in explaining their variability.[1]

Beaudry & Portier (2006) were the first to document using U.S. data that news shocks lead to positive comovement of consumption, hours worked, and investment, and account for the bulk of their variability at business-cycle frequencies. Beaudry & Lucke (2010) and Beaudry & Portier (2014) reach essentially the same conclusions. These findings have been challenged, however, by some scholars who questioned the underlying identification strategies.[2] Using an alternative, more flexible, identification approach, Barsky & Sims (2011) find that good news about future technology tend to raise consumption but to decrease output, hours worked, and investment in the short run.[3] They also find that inflation declines sharply and persistently in response to a positive realization of the news shock; a result deemed puzzling in light of the standard New Keynesian model.[4] Finally, though Barsky & Sims (2011) find that news shocks account for a significant fraction of output variability at business-cycle frequencies, they invoke the negative comovement to conclude that these shocks are unlikely to be a major driver of business cycles. These findings are confirmed by subsequent studies that propose alternative but related methodologies to Barsky & Sims' (e.g., Forni et al. (2014), Barsky et al. (2015), and Kurmann & Sims (2017)).

Existing empirical approaches to identify news shocks about future productivity are based on the premise that total factor productivity (TFP) is entirely and exclusively driven by two orthogonal disturbances: unanticipated and news shocks, the latter generally affecting

[1] An alternative approach to evaluate the importance of news shocks has been to estimate/calibrate dynamic stochastic general-equilibrium (DSGE) models that feature anticipated shocks to technology. This approach has been pursued by Jaimovich & Rebelo (2009), Fujiwara et al. (2011), Karnizova (2012), Schmitt-Grohé & Uribe (2012), and Khan & Tsoukalas (2012).

[2] Beaudry & Portier (2006), Beaudry & Lucke (2010), and Beaudry & Portier (2014) estimate small-scale systems (two to five equations) in which news shocks are identified using a mix of short- and long-run restrictions. Kurmann & Mertens (2014) show that Beaudry & Portier (2006)'s identification scheme does not have a unique solution when applied to a Vector Error Correction Model (VECM) with more than two variables. This identification scheme is therefore uninformative about the effects of news shocks and their importance for business cycles. Kurmann & Mertens (2014) further point out that the validity of the identification strategy proposed by Beaudry & Lucke (2010) critically depends on the plausibility of zero restrictions for other non-news shocks necessary to identify news shocks. Finally, Forni et al. (2014) argue that small-scale VARs and VECMs do not contain enough information to recover anticipated technology shocks from observable variables, a problem commonly known as *non-fundamentalness*.

[3] Barsky & Sims (2011) identify the news shock as the shock that best explains future movements in total factor productivity not accounted for by its own innovation.

[4] See, for instance, Jinnai (2013), Barsky et al. (2015), and Kurmann & Otrok (2014).

TFP with a lag. This assumption is consistent with the standard treatment of TFP in theoretical macroeconomic models. Hence, the above-mentioned studies invariably include a measure of TFP in the information set when attempting to identify news shocks from the data.

In this paper, we argue that the TFP measures typically utilized in the empirical literature contain important measurement errors that call into question the interpretation of TFP as a pure measure of technology. This is despite the corrections aiming at purging measured TFP of its non-technological component by controlling for unobserved variations in labor and capital. Most importantly, we demonstrate that the negative comovement of macroeconomic aggregates and the disinflation puzzle documented in recent empirical studies are spurious and are just an artifact of using a polluted measure of technology. In fact, we show that the news shocks identified in these studies are mostly picking up the effects of unanticipated technology shocks.

We document the severity of measurement errors in the adjusted TFP measure constructed by Fernald (2014) — which is the most widely used TFP series — by examining the dynamic effects of an unanticipated technology shock, identified as the reduced-form innovation to TFP, as is done in *all* existing VAR-based studies on news shocks.[5] The most revealing symptom of the presence of measurement errors is that unanticipated technological improvements are found to be inflationary, an outcome that runs against the conventional interpretation of surprise technology shocks as supply shocks, and violates the prediction of any sensible theory of aggregate fluctuations. A favorable surprise technology shock is also found to have counter-intuitive effects on stock prices and consumer confidence, which are initially unresponsive to the shock but fall persistently in the subsequent periods. We interpret these anomalous responses as an indication that the TFP series used in the empirical literature is an uncleansed measure of technology. Since a correct identification of news shocks hinges on the surprise technology shocks being properly identified, measurement errors in TFP are likely to undermine existing identification approaches.

We then propose an agnostic identification strategy that is robust to the presence of measurement errors in TFP. Our methodology relaxes the assumption that only technological shocks can affect measured TFP. Instead, we allow for the existence of non-technology shocks, which may capture measurement errors arising from the imperfect observability of inputs and their utilization rates, from the potential misspecification of the production function, and from aggregation bias. Non-technology shocks may affect measured TFP contemporaneously or at any future horizon, just like surprise technology shocks. To identify the latter, we rely on the sign-restriction approach proposed by Mountford & Uhlig (2009), imposing a negative sign on the inflation response to a positive shock. Hence, by construction, our strategy avoids the inflation anomaly engendered by identification schemes that associate surprise technology shocks with reduced-form innovations to TFP.

[5] The only exception is the study by Kurmann & Sims (2017), in which there is no attempt to identify surprise technology shocks.

We then extract the news shock as the linear combination of reduced-form innovations that is orthogonal to the surprise technology shock and that maximizes the contribution of the news shock to the forecast-error variance of TFP at a long but finite horizon. The argument underlying this criterion, originally proposed by Francis et al. (2014) and commonly referred to as the Max Share, is that the contribution of non-technology shocks to movements in TFP is likely to be negligible at very low frequencies.

We take our agnostic approach to the data by estimating a seven-variable VAR similar to that considered by Barsky & Sims (2011), first using their original data set, which spans the period 1960Q1–2007Q3, and then using an updated sample that extends the data coverage to 2016Q4. We find that non-technology shocks account for nearly half of the forecast error variance of Fernald's TFP series at the one-quarter horizon. This observation confirms the existence of non-trivial measurement errors in measured TFP and raises skepticism about available estimates of the effects of news shocks. Our results also show that the estimated effects of unanticipated technology shocks are remarkably consistent both with the predictions of the medium-scale New Keynesian model of Smets & Wouters (2007) and with the empirical evidence based on identification via long-run restrictions. In addition to being disinflationary by construction, an unanticipated technological improvement leads to a persistent and hump-shaped increase in consumption and output and to a short-term decline in hours worked. Moreover, the shock is found to have a positive effect on stock prices and consumer confidence.

Turning to the effects of news shocks, we find striking differences in the results across the two sample periods. In Barsky & Sims' original sample, we find little evidence of comovement (positive or negative) between consumption, output, and hours worked: while consumption increases following a favorable news shock, the initial response of output and hours worked is small and statistically indistinguishable from zero. Our results also indicate that the inflation response is muted and statistically insignificant at all horizons. In other words, the disinflation puzzle vanishes under our identification strategy. More generally, our estimated effects of news shocks are largely in line with the predictions of the Smets & Wouters (2007) model but differ markedly from those based on Barsky & Sims' approach. The latter turn out to be very similar to our estimated effects of a *surprise* technology shock, pointing to a misidentification of the news shock. Finally, variance-decomposition results and the historical decomposition of the time series of consumption, output, and hours strongly suggest that news shocks have not been a major contributor to business-cycle fluctuations in the sample ending in 2007.

A completely different portrait, however, is obtained from the updated sample. A favorable news shock identified using our agnostic strategy triggers a significant and persistent increase in consumption, output, and hours worked. Importantly, this simultaneous increase — indicative of positive comovement — occurs even before TFP starts to increase. Despite the fact that the disinflation puzzle essentially disappears under our empirical strategy, the estimated responses match rather poorly those implied by the Smets & Wouters (2007) model. We also find that news shocks account for roughly 40 to 60 percent of the forecast error variance of consumption, output, and hours worked at business-cycle frequencies, and that they explain a significant share of the decline in these quantities during the recent U.S. downturns, including the Great Recession. Together, these findings indicate that the importance of news shocks for business-cycle fluctuations has substantially increased in recent years, a conclusion that is inconsistent with the prediction based on Barsky & Sims' approach and with the results reached by Forni et al. (2014), Barsky et al. (2015), and Kurmann & Sims (2017).

The presumption that TFP is measured with error is of course not new; it has been discussed, for instance, in Christiano et al. (2004), Basu et al. (2006), and Fernald (2014). In a contemporaneous paper closely related to ours, Kurmann & Sims (2017) also study the implications of measurement errors in TFP for the identification of news shocks. These authors, however, do not establish a link between the anomalous responses to a surprise technology shock and the existence of measurement errors in TFP. Instead, their suspicion of the presence of such errors is based on the sensitivity of the estimated effects of news shocks using Barsky & Sims' methodology to revisions in Fernald's TFP series. Kurmann & Sims (2017) document that these revisions mainly reflect changes in the estimate of factor utilization, and argue that mis-measured utilization invalidates the identifying restriction that news shocks do not affect adjusted TFP on impact. Based on an identification strategy that relaxes this restriction and relies on the Max Share criterion to extract the news shock, they obtain very similar results to those documented by Barsky & Sims (2011) — namely, a negative comovement between consumption and hours and a limited contribution of news shocks to business-cycle fluctuations — with the difference that the results remain robust when the most recent data are used.

A crucial assumption of Kurmann & Sims' identification scheme is that the news shock is not orthogonalized with respect to the surprise technology shock (which is not identified). The two shocks are therefore likely to be muddled up since they both affect TFP in the short and in the long run, making it impossible — without further assumptions — to disentangle their respective contribution to the forecast error variance of TFP at any given horizon. Importantly, when we impose the orthogonality between the news and the surprise technology shock while relaxing the zero-impact restriction, we find no evidence of negative comovement and a significant role of news shocks in explaining aggregate fluctuations at business-cycle frequencies in the updated sample. In fact, our results are almost identical to those obtained by imposing the zero-impact restriction. This suggests that Kurmann & Sims' approach may be confounding news and surprise technology shocks.

The rest of this paper is organized as follows. Section 3.2 discusses the symptoms of measurement errors in TFP. Section 3.3 presents our agnostic identification strategy. Section 3.4 discusses the results based on Barsky & Sims (2011) original data and on an updated sample. Section 3.5 studies the robustness of our results when we relax the zero-impact restriction. Section 3.6 concludes.

3.2 The Inflation Anomaly and Other Symptoms of Measurement Errors in TFP

In this section, we illustrate the extent to which the effects of unanticipated technology shocks typically reported in the VAR-based "news" literature are inconsistent with the predictions of New Keynesian models and, for that matter, any sensible theory of aggregate fluctuations. We view these inconsistencies as symptoms of the presence of measurement errors in the TFP series commonly used in the literature.

3.2.1 Unanticipated technology shocks: measurement...

In the VAR-based literature on news shocks, unanticipated technology shocks are usually identified as the reduced-form innovations to TFP. Formally, let y_t be a $k \times 1$ vector of observables of length T, which includes TFP and which has the following moving-average (MA) representation

$$y_t = B(L)u_t,$$

where u_t is a $k \times 1$ vector of statistical innovations, whose variance-covariance matrix is denoted by Σ. Let ϵ_t be a $k \times 1$ vector of structural innovations, including the unanticipated technology shock, whose variance-covariance matrix is normalized to I_k. If a linear mapping between the statistical innovations, u_t, and the structural shocks, ϵ_t, exists, then we can write

$$u_t = A\epsilon_t,$$

where the impact matrix, A, must be such that $AA' = \Sigma$. Assuming (without loss of generality) that TFP is ordered first in y_t and that the unanticipated technology shock is ordered first in ϵ_t, a Cholesky decomposition of Σ ensures that the surprise technology shock is proportional to the statistical innovation to TFP.

We use the strategy above to measure the effects of a surprise technology shock within a seven-variable VAR similar to that estimated by Barsky & Sims (2011). The vector of observables includes adjusted TFP, output, consumption, hours, inflation, stock prices, and consumer confidence, measured at a quarterly frequency. We start by using Barsky and Sims' original data, which span the period 1960Q1–2007Q3; we then update the sample by extending it to 2016Q4.[6]

The results are shown with solid black lines in Figure 3.1.[7] The (one-standard-error) confidence intervals around the estimated impulse responses are computed using the bias-

[6] The series used in estimation are constructed as follows. Adjusted TFP is the quarterly series constructed by Fernald (2014), which controls for unobserved input variation. Output is measured by the log of real GDP in the non-farm business sector. Consumption is measured by the log of real personal spending on non-durables and services. Hours are measured by the log of total hours worked in the non-farm business sector. Output, consumption and hours are expressed in per capita terms by dividing them by the civilian, noninstitutional population, age 16 and over. Inflation is measured by the percentage change in the CPI for all urban consumers. Stock prices are measured by the log of the S&P index. Consumer confidence is retrieved from the Michigan Survey of Consumers.

[7] These results are based on a VAR with 3 lags. Alternative lag lengths yield similar results.

corrected bootstrap procedure proposed by Kilian (1998). A surprise technology shock triggers a transitory increase in TFP, output, and consumption. In all three cases, the estimated response is rather monotonic and the variable reverts to its pre-shock level rather rapidly. In contrast, hours worked exhibit a relatively muted — and mostly statistically insignificant — response. The Figure also shows that, in response to the identified surprise technology shock, inflation rises persistently and in a hump-shaped manner, with a peak occurring at around 10 quarters after the shock. Stock prices and consumer confidence, in contrast, are unresponsive on impact and eventually fall below their pre-shock levels for a prolonged period of time. Very similar results are reported by Forni et al. (2014), Barsky et al. (2015), Fève & Guay (2016), and Kurmann & Sims (2017).

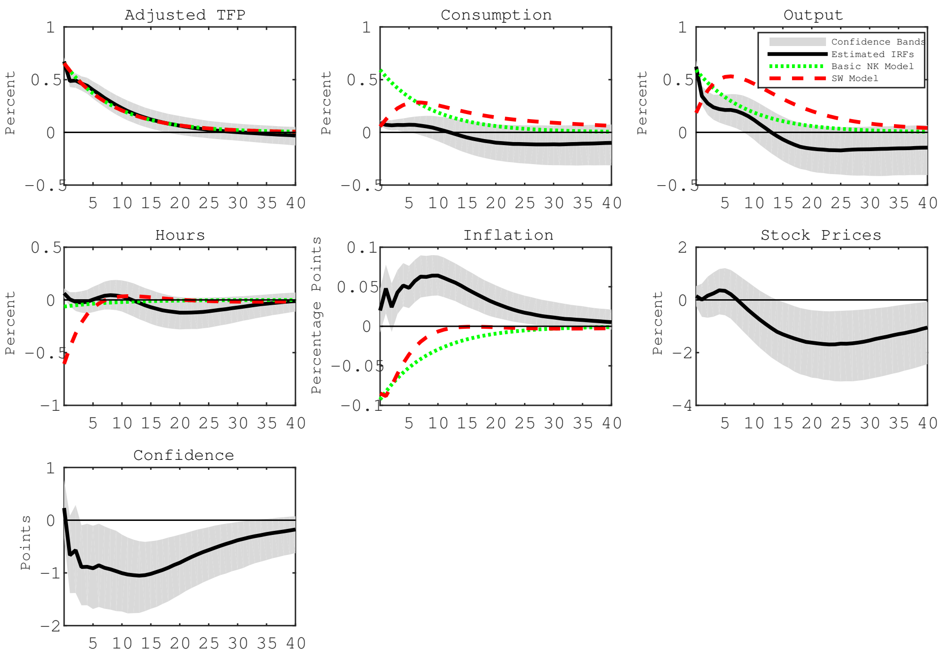

Figure 3.1: Impulse responses to a surprise technology shock. Sample: 1960Q1–2007Q3.

Notes: The figure shows the impulse responses to a surprise technology shock. The solid lines are the median impulse responses estimated based on the reduced-form innovation to TFP. The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications. The dotted lines are the impulse responses obtained from the standard New Keynesian model. The dashed lines are the impulse responses obtained from the Smets & Wouters (2007) model.

When we extend the sample to 2016Q4, two notable differences with respect to the results above stand out (see Figure 3.2). First, hours worked now fall initially in response to the shock, but their response remains mostly statistically insignificant. Second, stock prices and consumer confidence now rise for about three quarters after the shock, but they continue to decline persistently during the subsequent quarters. These two exceptions aside, the results based on the updated sample are very similar to the original ones. In particular, the response of TFP, output and consumption are transitory, inflation rises persistently and in a hump-shaped manner, and consumer confidence falls persistently with a delay.

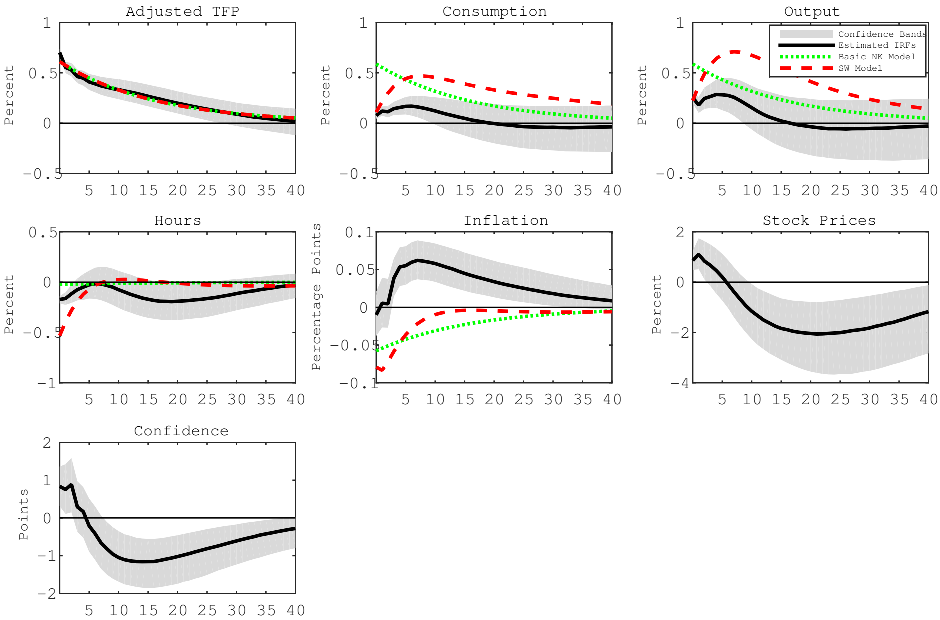

Figure 3.2: Impulse responses to a surprise technology shock. Sample: 1960Q1–2016Q4.

Notes: The figure shows the impulse responses to a surprise technology shock. The solid lines are the median impulse responses estimated based on the reduced-form innovation to TFP. The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications. The dotted lines are the impulse responses obtained from the standard New Keynesian model. The dashed lines are the impulse responses obtained from the Smets & Wouters (2007) model.

3.2.2 ... and theory

How do the empirical findings discussed in the previous section compare with the predictions of New Keynesian theory of aggregate fluctuations? We answer this question by studying the effects of unanticipated technology shocks both within the simplest version of the New Keynesian model and the more realistic medium-scale version proposed by Smets & Wouters (2007). To do so, we assume that the log of TFP (in deviation from its mean), a_t, is governed by the following process:

$$a_t = \rho_a a_{t-1} + x_{t-1} + \epsilon_t^s, \tag{3.1}$$

$$x_t = \rho_x x_{t-1} + \epsilon_t^n, \tag{3.2}$$

where ϵ_t^s and ϵ_t^n are, respectively, the surprise and anticipated (or news) technology shocks, and $0 \leq \rho_a, \rho_x < 1$. Notice that ρ_x is irrelevant to the dynamic effects of the surprise shock and thus ρ_a and the size of the disturbance ϵ_t^s are the only parameters that one needs to calibrate to study those effects. We choose those two parameters such that the implied response of TFP to the surprise technology shock mimics as closely as possible the response estimated from the data. The model-based responses of TFP, consumption, output, hours, and inflation are superimposed on their empirical counterparts in Figures 3.1 and 3.2.

The basic New Keynesian model

Consider first the basic New Keynesian model, summarized by the following log-linearized equations (around a zero-inflation steady state):[8]

$$
\begin{aligned}
c_t &= y_t, && (3.3)\\
y_t &= a_t + n_t, && (3.4)\\
mc_t &= \sigma c_t + +\varphi n_t - a_t, && (3.5)\\
c_t &= E_t c_{t+1} - \sigma^{-1}(i_t - E_t \pi_{t+1} - \ln\beta), && (3.6)\\
\pi_t &= \beta E_t \pi_{t+1} + \lambda mc_t, && (3.7)\\
i_t &= \ln\beta + \phi_\pi \pi_t + \phi_y (y_t - y_t^f), && (3.8)
\end{aligned}
$$

where c_t is consumption, y_t is output, n_t is hours worked, mc_t is real marginal cost, π_t is the inflation rate, i_t is the nominal interest rate, and $y_t^f = (1+\varphi)(\sigma+\varphi)^{-1} a_t$ is the flexible-price (or natural) level of output. All the variables are expressed as percentage deviations from their steady-state values except π_t and i_t, which are expressed in levels. The parameters are defined as follows: $\sigma > 0$ is the inverse of the elasticity of intertemporal substitution, $\varphi > 0$ is the inverse of the Frisch elasticity of labor supply, $0 < \beta < 1$ is the discount factor, $\lambda = (1-\theta)(1-\beta\theta)/\theta > 0$, $0 < \theta < 1$ is the Calvo probability of not changing prices, and $\phi_\pi, \phi_y > 0$ are the coefficients attached to inflation and the output gap in the interest rate rule.

Model (3.3)–(3.8) can be solved analytically to determine the effects of a surprise technology shock. Assuming that $\epsilon_t^n = 0$ for all t, one can use the method of undetermined coefficients to show that

$$\pi_t = -\frac{\sigma\lambda(1+\varphi)(1-\rho_a)}{\Delta_a} a_t,$$

where $\Delta_a = \lambda(\sigma+\varphi)(\phi_\pi - \rho_a) + (1-\beta\rho_a)[\sigma(1-\rho_a) + \phi_y] > 0$.[9] Since the numerator in the expression above is positive, an unanticipated technological improvement will cause inflation to fall persistently as long as $\rho_a < 1$. This disinflationary effect reflects the persistent fall in real marginal cost or, equivalently, the negative output gap resulting

[8] This is essentially the model presented in Gali (2008). The only difference is that we assume (for simplicity) constant returns to scale in the production technology. This simplification has no impact on the results.

[9] The necessary and sufficient condition for the existence of a unique linear rational expectations equilibrium is given by $\lambda(\sigma+\varphi)(\phi_\pi - 1) + (1-\beta)\phi_y > 0$. It is straightforward to see that this condition implies that $\Delta_a > 0$.

from the shock.[10] This can be seen by noticing that

$$mc_t = -\frac{\sigma\left(1+\varphi\right)\left(1-\rho_a\right)\left(1-\beta\rho_a\right)}{\Delta_a}a_t.$$

The surprise technology shock has a positive effect on output (and thus consumption) but an ambiguous effect on hours worked. The solutions for these variables are given by

$$\begin{aligned} y_t &= \frac{\left(1+\varphi\right)\left[\lambda\left(\phi_\pi-\rho_a\right)+\left(\sigma+\varphi\right)^{-1}\left(1-\beta\rho_a\right)\phi_y\right]}{\Delta_a}a_t, \\ n_t &= \left\{\frac{\left(1+\varphi\right)\left[\lambda\left(\phi_\pi-\rho_a\right)+\left(\sigma+\varphi\right)^{-1}\left(1-\beta\rho_a\right)\phi_y\right]}{\Delta_a}-1\right\}a_t. \end{aligned}$$

Under plausible parameter values, however, hours worked fall in response to a positive unanticipated technology shock. The responses depicted in Figures 3.1 and 3.2 (with green dotted lines) are obtained using the following standard parameterization of the model: $\sigma = \varphi = 1$, $\beta = 0.99$, $\theta = 0.75, \phi_\pi = 1.5, \phi_y = 0.125$. Under these parameter values, a positive surprise shock to technology raises output and consumption and decreases hours worked and inflation.

The dynamic responses implied by the model hardly match those estimated from the data, but the most striking discrepancy concerns the response of inflation, which has the opposite sign and a completely different shape relative to what is predicted by the VAR.

The Smets and Wouters (2007) model

Next, consider the medium-scale model developed by Smets & Wouters (2007). To conserve space, we only summarize the main features of the model and refer the reader to their paper for a more detailed description. The model features a representative household whose preferences exhibit habit formation in consumption. The final good is produced using an aggregator of intermediate goods that exhibits a non-constant elasticity of substitution. Intermediate goods are produced using a technology that depends on TFP, labor, and capital, and that exhibits variable capital utilization and fixed costs. Capital accumulation is subject to investment adjustment costs. Both prices and wages are set in a staggered fashion à la Calvo, whereby the non-optimizing agents partially index their prices and wages to past inflation, thus giving rise to a New Keynesian Phillips curve that depends not only on current and expected future inflation but also past inflation. Monetary policy follows an interest rate rule with a smoothing component. The model is estimated by Bayesian techniques using U.S. data over the period 1966Q1–2004Q4.

We use Smets and Wouters' posterior means for the structural parameters to generate the implied responses to an unanticipated positive technology shock, which are represented by the dashed red lines in Figures 3.1 and 3.2. Despite some quantitative differences, these

[10] By iterating equation (3.7) forward, inflation can be expressed as a discounted sum of current and expected future real marginal costs.

responses are in line with the predictions of the basic New Keynesian model: output and consumption rise while hours worked and inflation fall in response to the shock. The fall in inflation persists for about eight quarters after the shock, which is in stark contrast with the positive response obtained from the VAR.[11] Notice also that the VAR-based responses of output and consumption lack the persistent and hump-shaped pattern implied by the model.

3.2.3 Discussion

As we have just shown, reduced-form innovation to TFP are found to be inflationary, an outcome that runs against the conventional interpretation of technology shocks as supply shocks, and contradicts the prediction of any sensible macroeconomic model. It is also at odds with the results reported by a number of empirical studies that rely on the long-run restriction approach proposed by Gali (1999) to identify exogenous technology shocks (e.g., Edge et al. (2003), Christiano et al. (2003), Feve & Guay (2010)). Moreover, the result that technology shocks have a delayed negative effect on stock prices and consumer confidence also appears hard to reconcile with the view that technology enhances efficiency and raises the productive capacity of the economy.

These observations cast serious doubt on the interpretation of reduced-form innovations to TFP as pure unanticipated technological improvements. The identified shocks appear to be contaminated by other non-technological disturbances that also affect measured TFP contemporaneously and whose effects are akin to those of a demand shock. Since a proper identification of news shocks about future productivity hinges on purging TFP of its non-technological component, the anomalous responses just discussed suggest that existing methodologies — albeit sound in theory — may still fail to correctly identify news shocks and their effects due to measurement errors in TFP.

In the models discussed in Section 3.2.2, TFP is assumed to be exogenous to the state of the economy and, as such, is not expected to be affected by demand shocks — note that this is precisely the identifying assumption underlying the empirical literature on news shocks. TFP, however, is not readily observable in the data and must be inferred from production and input use, a task that poses a number of measurement challenges. First, some inputs may not be observable or measurable; second, input utilization varies in response to non-technology shocks; third, the production technology may have non-constant returns to scale; fourth, aggregating inputs across heterogeneous production sectors may introduce a bias. Failing to eliminate any of these potential sources of measurement errors may result in an incorrect measure of TFP and thus a poor proxy for technology. In their seminal paper, Basu et al. (2006) went a long way towards constructing a purified annual measure of technology by adjusting TFP for observed and unobserved input variations and non-constant returns to scale. The quarterly TFP series used in the empirical lit-

[11] A persistent decline in inflation following a favorable surprise technology shock is also predicted by the New Keynesian models estimated by Ireland (2004) and Altig et al. (2011), though the inflation response is relatively small in magnitude in the latter case.

erature on news shocks was constructed by Fernald (2014) following Basu et al. (2006)'s methodology but without correction for non-constant returns to scale since the industry level data needed for this correction are only available at an annual frequency.
To get a sense of how this impacts the measurement of TFP, we plot in Figure 3.3 annual TFP growth based on the measures constructed by Fernald (2014) and Basu et al. (2006) for the period 1960–1996.[12] Although there is some similarity between the two series, their correlation is modest (0.57), suggesting that the constant-returns-to-scale assumption underlying the construction of the quarterly TFP series is counterfactual and is likely to be one of the culprits for the anomalous responses documented above.

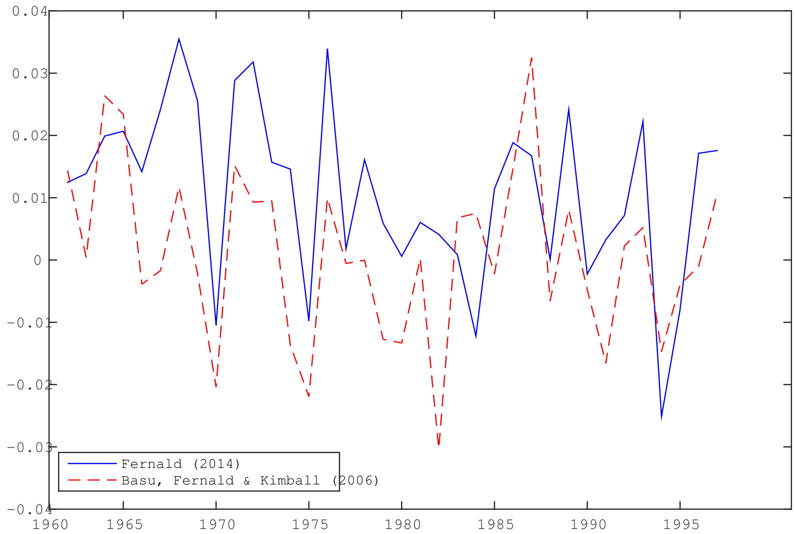

Figure 3.3: Annual TFP growth based on the series constructed by Fernald (2014)'s and Basu et al. (2006)'s.

To further illustrate the importance of this assumption as a potential source of measurement errors, we estimate the effects of a surprise technology shock identified as the reduced form innovation to Basu et al. (2006)'s series using the same observable variables as in section 3.2.1, measured annually. The estimated impulse responses and their confidence bands are shown in Figure 3.4, in which a period corresponds to a year.[13] The figure shows that, following a positive technology shock, output remains essentially unresponsive on impact but increases in a hump-shaped manner during the subsequent years, whereas hours worked fall significantly at the time of the shock. Inflation also falls sharply on impact, consistently with the expected disinflationary effect of a technological improvement, and in sharp contrast with the rise in inflation obtained using the quarterly TFP series. This observation hints at the fact that Basu et al. (2006)'s TFP series is less

[12] Basu et al. (2006)'s TFP series ends in 1996.

[13] The results reported in Figure 3.4 are based on a VAR with one lag. We obtain very similar results when we include two lags. Because we are estimating a VAR with 7 variables using 36 annual observations, including more lags leaves too few degrees of freedom to obtain reliable estimates.

polluted by non-technological factors than Fernald (2014)'s quarterly series.

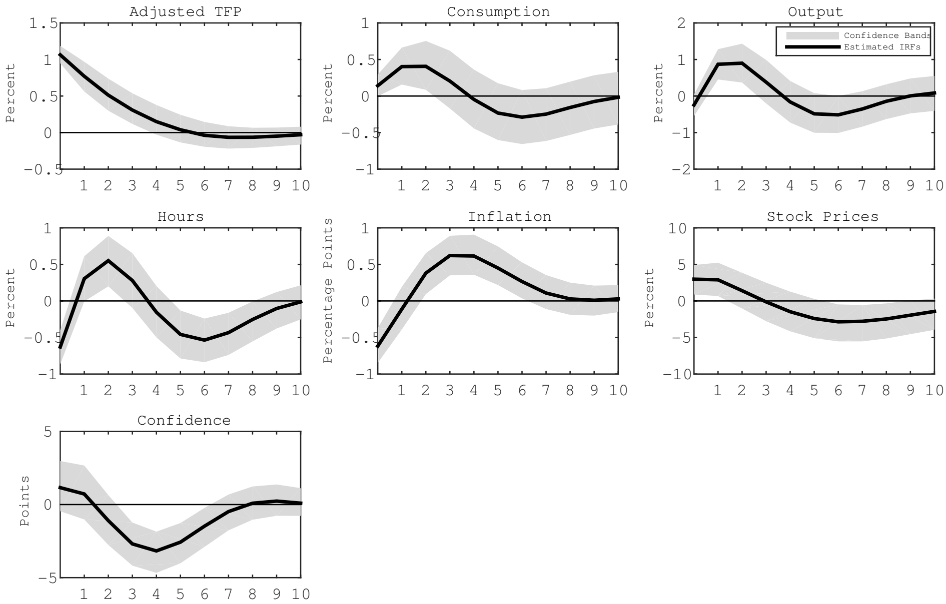

Figure 3.4: Impulse responses to a surprise technology shock based on Basu et al. (2006)'s annual TFP series.

Notes: The figure shows the impulse responses to a surprise technology shock. The solid lines are the median impulse responses estimated based on the reduced-form innovation to TFP. The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications.

Yet, Figure 3.4 shows that even Basu et al. (2006)'s purified TFP measure generates some anomalies that are hard to reconcile with conventional wisdom about the effects of technology shocks. For instance, the initial disinflationary effect of the shock is followed by a protracted episode (of several years) during which inflation is above average. Moreover, while stock prices initially rise in response to a positive technology shock, they decline persistently during the subsequent years. Likewise, the shock triggers a delayed fall in consumer confidence that persists for a prolonged period of time. These responses cast doubt on the interpretation of the shock as a pure technological disturbance.

In sum, despite the colossal work carried out by Basu et al. (2006) and Fernald (2014) to construct a cleansed measure of technology, it is probably unrealistic to believe that the corrected TFP series is purged of all its non-technological factors, which in turn suggests that TFP-based measures of technology shocks will most likely be contaminated by measurement errors. This conclusion motivates the agnostic approach that we describe in the next section.

3.3 An Agnostic Identification Approach

3.3.1 Idea

The maintained assumption underlying the empirical identification of news shocks about future productivity is that measured TFP is exclusively driven by surprise and anticipated technology shocks, the latter affecting TFP only with a lag. The common approach to identify the news shock is then to select the linear combination of reduced-form innovations that best explains (or forecasts) future movement in TFP while being orthogonal to the surprise technology shock. This strategy will correctly identify news shocks only to the extent that surprise technology shocks are the only disturbances that affect measured TFP contemporaneously, which, as we just argued above, seems highly unlikely.

We propose an alternative empirical strategy based on the assumption that measured TFP is affected by two types of disturbances: technological and non-technological shocks. The latter capture measurement errors due to the imperfect observability of inputs and their utilization rates, to the potential misspecification of the production function, and to aggregation bias. From this perspective, it may be inappropriate to characterize these shocks as structural, given that they do not bear a clear economic interpretation. However, this is not a concern for our methodology since we need not identify these shocks; we simply allow them to affect measured TFP contemporaneously and at any future horizon, just as surprise technology shocks.

To identify the surprise technology shock, we adopt an agnostic strategy based on the sign-restriction approach proposed by Mountford & Uhlig (2009). More specifically, we select the impulse vector that (most markedly) satisfies the restriction that inflation falls for at least eight quarters after the shock, consistently with the prediction of the Smets & Wouters (2007) model. Hence, by construction, our strategy avoids the inflation anomaly engendered by identification schemes that associate surprise technology shocks with TFP innovations. We then identify the news shock as the linear combination of reduced-form innovations that is orthogonal to the surprise technology shock and that maximizes the contribution of the news shock to the forecast-error variance of TFP at a long but finite horizon, H. The latter criterion, initially proposed by Francis et al. (2014) and commonly referred to as the Max Share, differs from the one used by Barsky & Sims (2011), which involves maximizing the contribution of the news shocks to the forecast error variance of TFP over all horizons up to a finite truncation horizon. Barsky & Sims' approach has been criticized on the ground that it may confound shocks that have either permanent or temporary effects on TFP, and has been shown to be quite sensitive to the truncation horizon (see Beaudry et al. (2011)). Since our approach allows for the presence of non-technology shocks, whose effects on measured TFP are likely to be much more important at short horizons than at more distant ones, this makes the case for using the Max Share even stronger.

3.3.2 Implementation

Let $\tilde{A}$ denote the Cholesky decomposition of Σ and assume again that TFP is ordered first in y_t. Any impact matrix $A_0 = \tilde{A}D$, where D is an orthonormal matrix, also satisfies the requirement $A_0 A'_0 = \Sigma$. Let γ_j denote the jth column of D, ϵ_1 denote the surprise technology shock, and ϵ_2 denote the news shock.

We identify the surprise technology shock by selecting the orthonormal matrix D that satisfies the requirement that inflation does not increase during the first eight quarters after the shock while yielding the largest response in the desired direction. Because the impulse vector to this shock is $\tilde{A}\gamma_1$ (the first column of $\tilde{A}D$), we only need to characterize γ_1.

Denote by $r_{j,i}(h)$ the impulse response of the jth variable to the ith column of $\tilde{A}$ at horizon h (that is, the reduced-form impulse response), and by $r_i(h)$ the $k-$dimensional column vector $[r_{1,i}(h), \cdots, r_{k,i}(h)]$. The $k-$dimensional impulse response $r_{\gamma_1}(h)$ to the impulse vector $\tilde{A}\gamma_1$ is given by

$$r_{\gamma_1}(h) = \sum_{i=1}^{k} \gamma_{i,1} r_i(h),$$

where $\gamma_{i,1}$ is the ith entry of γ_1.

Following Mountford & Uhlig (2009)'s approach, we select the vector γ_1 of unit length that solves the following minimization problem:

$$\min_{\{\gamma_1\}} \Psi(\tilde{A}\gamma_1),$$

with the criterion function, $\Psi(\tilde{A}\gamma_1)$, being given by

$$\Psi(\tilde{A}\gamma_1) \equiv \sum_{h=0}^{7} f\left(\frac{r_{\pi,\gamma_1}(h)}{s_\pi}\right),$$

where the loss function, f, is such that $f(x) = 100x$ if $x > 0$ and $f(x) = x$ if $x \leq 0$, and s_π is the standard deviation of the reduced-form innovation to inflation. The criterion $\Psi(\tilde{A}\gamma_1)$ therefore strongly penalizes impulse vectors that generate a positive inflation response at any given horizon. If multiple impulse vectors are consistent with the imposed sign restriction on the response of inflation, the unique solution to the minimization problem above will be the impulse vector that yields the largest fall in inflation over eight quarters.

Once the surprise technology shock, ϵ_1, is identified, we identify the news shock, ϵ_2, as the linear combination of the reduced-form residuals that is orthogonal to ϵ_1 and that explains the largest fraction of the forecast error variance of TFP at a long but finite horizon, H. The h-step ahead forecast error of vector y is

$$y_{t+h} - E_t y_{t+h} = \sum_{\tau=0}^{h-1} B_\tau \tilde{A} D \epsilon_{t+h-\tau}.$$

Denoting by $\Omega_{i,j}(h)$ the share of the forecast error variance of variable i attributable to structural shock j at horizon h, this quantity is given by

$$\Omega_{i,j}(h) \equiv \frac{e_i^{'}\left(\sum_{\tau=0}^{h-1} B_\tau \tilde{A} D e_j e_j^{'} D^{'} \tilde{A} B_\tau^{'}\right) e_i}{e_i^{'}\left(\sum_{\tau=0}^{h-1} B_\tau \Sigma B_\tau^{'}\right) e_i} = \frac{\sum_{\tau=0}^{h-1} B_{i,\tau} \tilde{A} \gamma_j \gamma_j^{'} \tilde{A} B_{i,\tau}^{'}}{\sum_{\tau=0}^{h-1} B_{i,\tau} \Sigma B_{i,\tau}^{'}},$$

where

$$B_{i,\tau} = e_i^{'} B_\tau, \qquad \gamma_j = D e_j,$$

and e_i is a selection vector with 1 in the ith position and zero elsewhere. The identification of the news shock therefore amounts to selecting the vector γ_2 that solves the following maximization problem:

$$\max_{\{\gamma_2\}} \Omega_{1,2}(H) \equiv \frac{\sum_{\tau=0}^{H-1} B_{1,\tau} \tilde{A} \gamma_2 \gamma_2^{'} \tilde{A} B_{1,\tau}^{'}}{\sum_{\tau=0}^{H-1} B_{1,\tau} \Sigma B_{1,\tau}^{'}}$$

s.t.

$$\gamma_2(1) = 0, \qquad \gamma_2^{'}\gamma_1 = 0, \qquad \gamma_2^{'}\gamma_2 = 1.$$

The first constraint ensures that the news shock does not affect TFP contemporaneously; the second constraint ensures that the news shock is orthogonal to ϵ_1; and the third constraint ensures that γ_2 is a column vector of an orthonormal matrix. In practice, we choose $H = 80$ quarters.

3.4 Results

We apply our agnostic identification strategy to the same seven-variable VAR estimated by Barsky & Sims (2011). We consider two data sets: the one originally used by these authors, which spans the period 1960Q1–2007Q3, and an updated data set that extends the data coverage through 2016Q4. For each of these samples, we discuss the impulse responses to a surprise and an anticipated technology shock, the contribution of news shocks to the forecast error variance of macroeconomic aggregates, and their historical decomposition. In the process, we contrast our findings with those obtained using Barsky & Sims' methodology.

3.4.1 Sample period 1960Q1–2007Q3

Impulse responses We start by discussing the estimated impulse responses to a surprise and an anticipated technology shock. To gauge these responses from the standpoint of New Keynesian theory, we compare them with those implied by the Smets & Wouters (2007) model. To do so, we again assume that TFP is described by process (3.1)–(3.2) and calibrate the parameters ρ_a and ρ_x and the size of the disturbances ϵ_t^s and ϵ_t^n so as to replicate as closely as possible the estimated response of TFP to the surprise and the news shock. The confidence intervals around the estimated impulse responses are computed using Kilian (1998)'s bias-corrected bootstrap procedure.

The estimated impulse responses to a surprise technology shock are reported in the right column of Figure 3.5. For ease of comparison with the results based on reduced-form innovations to TFP (as in Barsky & Sims (2011) and the rest of the empirical literature on news shocks), the left column of Figure 3.5 reproduces the responses reported in Figure 3.1 using the same scale for each response as in the right column.

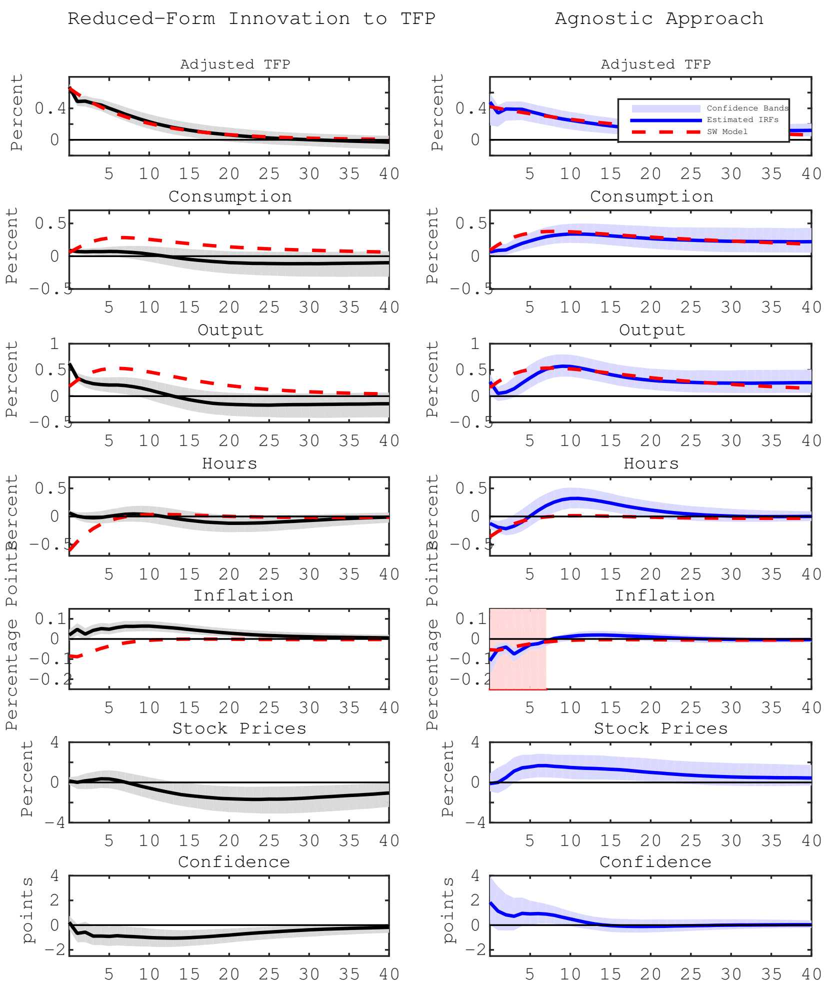

Figure 3.5: Impulse responses to a surprise technology shock. Sample: 1960Q1–2007Q3.

Notes: The figure shows the impulse responses to a surprise technology shock. The solid lines are the median impulse responses estimated based on the reduced-form innovation to TFP (left panels) and on the agnostic approach (right panels). The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications. The shaded red area indicates the horizons at which the inflation response is constrained to be negative. The dashed lines are the impulse responses obtained from the Smets & Wouters (2007) model.

TFP increases on impact and remains persistently higher than its pre-shock level, a pattern that contrasts with the rapid return obtained when surprise technology shocks are identified as TFP innovations (shown in the upper left panel of Figure 3.5).[14] Consumption and output also increase persistently and in a hump-shaped fashion. The estimated responses are remarkably similar to those implied by the Smets & Wouters (2007) model (particularly for consumption), and sharply contrast with the small, transitory and rather monotonic pattern obtained from the identification scheme associating the shock with the TFP innovation.

Hours worked initially fall for about five quarters, then increase in a hump-shaped manner before converging to their pre-shock level from above. This pattern is consistent with the prediction of the Smets & Wouters (2007) model, at least qualitatively, and differs from the muted reaction shown in the corresponding left panel. The result that unanticipated technological improvement has a contractionary effect on employment in the short run has been documented in several studies using different empirical approaches.[15]

Our estimated response for inflation is, by construction, restricted to be negative for the first eight quarters after the shock, as indicated by the shaded red area. Beyond that horizon, the inflation response becomes small and statistically insignificant. Interestingly, although our identification strategy does not impose a precise numerical value for the inflation response, the estimated response is strikingly similar to that implied by the Smets & Wouters (2007) model. The latter lies within the estimated confidence band at almost any given horizon.

Our identified surprise technology shock raises stock prices and consumer confidence. Stock prices are initially unresponsive but increase significantly and persistently during the subsequent quarters. The increase in consumer confidence is more transitory and is only statistically significant on impact and between the sixth and eighth quarters after the shock. These responses are at variance with the persistent decline in stock prices and consumer confidence shown in the left panels of Figure 3.5.

In sum, these findings show that identifying surprise technology shocks by restricting their effect on inflation to be negative produces impulse responses that are more consistent with conventional wisdom and better grounded in theory than those obtained by using reduced-form innovations to TFP as a measure of surprise technology shocks. Interestingly, our estimated responses mimic remarkably well those implied by the Smets & Wouters (2007) model. The latter mostly lie within the confidence bands of the VAR-based responses.

The estimated responses to a news shock are illustrated in the right column of Figure 3.6. The response of TFP is similar in shape but significantly smaller in magnitude than that based on Barsky & Sims' approach. An important conclusion from Barsky & Sims' paper is that output and hours worked initially decline in response to a favorable news shock about future productivity (see the third and fourth panels on the left column of Figure 3.6), an outcome that violates the predictions of the Smets & Wouters (2007) model. Both variables then rise persistently during the subsequent quarters, although the

[14] This is reflected in the larger estimate of the parameter ρ_a implied by our estimated response of TFP (0.956) than that implied by the TFP response estimated using Barsky and Sims' methodology (0.897).

[15] See Galí & Rabanal (2005) for a survey.

rise in hours is mostly statistically insignificant. A similar pattern for hours is reported by Forni et al. (2014), Barsky et al. (2015), and Kurmann & Sims (2017).[16] The short-run contractionary effect of the news shock on aggregate output and hours worked no longer occurs, however, when we use our agnostic empirical methodology, as the output response is now statistically insignificant during the first two quarters after the shock, and that of hours worked is statistically indistinguishable from zero at any given horizon. In other words, we find no evidence of comovement, either negative or positive, between macroeconomic aggregates conditional on our identified news shock.

Turning to the response of inflation, Barsky & Sims' approach implies that a favorable news shock about future technology decreases inflation sharply and persistently. This disinflationary effect, also documented by Forni et al. (2014), Barsky et al. (2015), Fève & Guay (2016), and Kurmann & Sims (2017), is puzzling in light of New Keynesian theory, as pointed out by Barsky & Sims (2009), Jinnai (2013), and Kurmann & Otrok (2014). In the context of the basic New Keynesian model presented in Section 3.2.2, it is possible to show (using the method of undetermined coefficients) that the initial response of inflation to a news shock is given by

$$\frac{d\pi_t}{d\epsilon_t^n} = \frac{\sigma\lambda(1+\varphi)\left[\lambda(\sigma+\varphi)(\phi_\pi-1)+(1-\beta)\phi_y-\beta\sigma(1-\rho_a)(1-\rho_x)\right]}{\Delta_a\Delta_x},$$

where $\Delta_x = \lambda(\sigma+\varphi)\left(\phi_\pi-\rho_x\right)+(1-\beta\rho_x)\left[\sigma\left(1-\rho_x\right)+\phi_y\right] > 0$. While the sign of the expression above is, in principle, ambiguous, it typically tends to be positive under sufficiently high values of ρ_a and ρ_x and a plausible calibration of the remaining parameters. Using the estimated values of ρ_a and ρ_x and the calibration discussed in Section 3.2.2, the basic New Keynesian model predicts a positive response of inflation to a favorable TFP news shock. The Smets & Wouters (2007) model also implies that inflation rises temporarily after a positive news shock but the response is tiny and essentially indistinguishable from 0 at any given horizon. This disinflation puzzle has prompted some researchers to suggest modifications to the prototype New Keynesian model so as to reconcile its predictions with the empirical evidence.[17] Contrasting with the existing evidence, however, our results indicate that the inflation response to a favorable news shock is rather muted and statistically insignificant at all horizons, consistently with the theoretical prediction. In other words, the disinflation puzzle vanishes under our agnostic identification strategy. The disinflationary effect documented in earlier studies appears to be an artifact of the misidentification of anticipated technology shocks, due to measurement errors in TFP.

[16] Forni et al. (2014)'s approach is based on an estimated factor-augmented VAR in which the news shock is identified as the shock that best anticipates TFP at the 60-quarter horizon while being orthogonal to the reduced-form innovation in TFP. Barsky et al. (2015) identify the news shock as the innovation in the expectation of TFP at a fixed horizon in the future (20 quarters). Kurmann & Sims (2017) rely on the Max Share method (with $H = 80$) but without imposing the orthogonality of the news shock with respect to current TFP.

[17] See, for instance, Jinnai (2013), Barsky et al. (2015), and Kurmann & Otrok (2014).

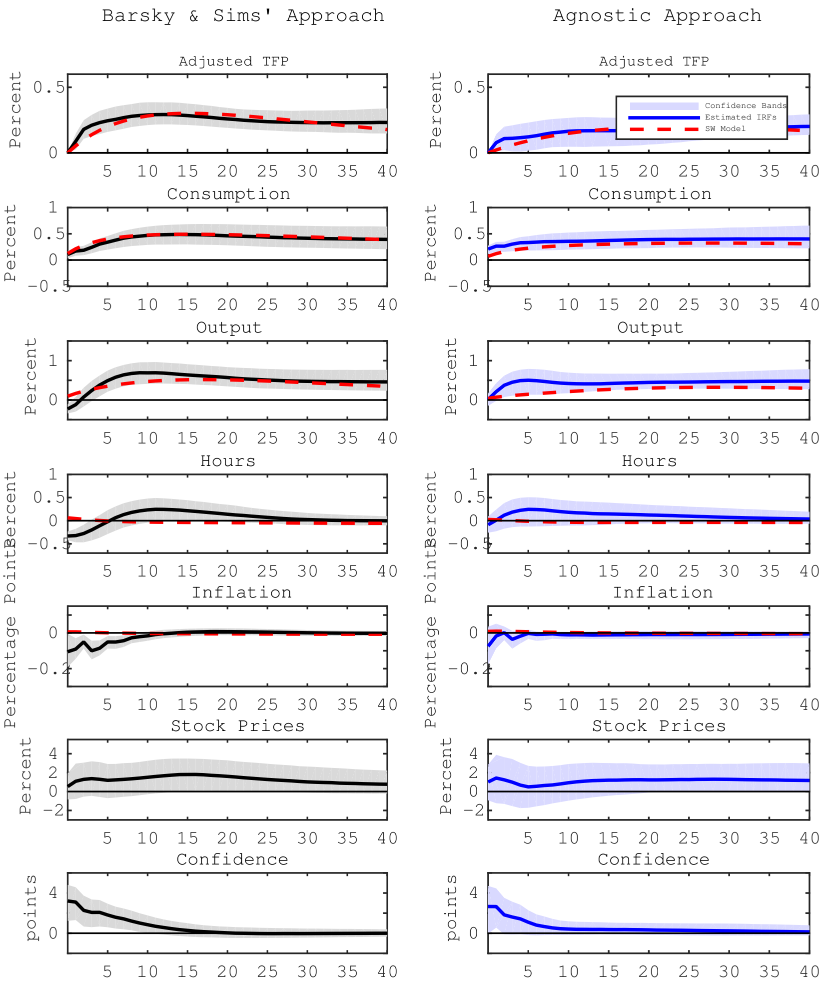

Figure 3.6: Impulse responses to a news shock. Sample: 1960Q1–2007Q3.

Notes: The figure shows the impulse responses to a news shock. The solid lines are the median impulse responses estimated based on Barsky and Sims' approach (left panels) and on the agnostic approach (right panels). The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications. The dashed lines are the impulse responses obtained from the Smets & Wouters (2007) model.

Variance decomposition Before evaluating the contribution of news shocks to the variability of macroeconomic variables, it is worth discussing the relative importance of the identified surprise technology shocks in explaining TFP. The results are reported in Table 3.1.[18] By construction, when surprise technology shocks are identified as the reduced-form innovations to TFP, they explain *all* of the forecast error variance of TFP at $h = 1$ (recall that the news shock does not affect TFP contemporaneously). Under our agnostic strategy, however, this need not be the case. In fact, our identified surprise technology shocks account for roughly half of the one-quarter ahead forecast error variance of TFP, thus implying that non-technological shocks (potentially reflecting measurement errors) account for the remaining half, which in turn raises a serious objection against the interpretation of the estimated TFP series as a purified measure of technology.

Table 3.1: Share of Forecast Error Variance of TFP attributed to Surprise Technology Shocks. Sample: 1960Q1–2007Q3.

	Horizon					
	$h = 1$	$h = 4$	$h = 8$	$h = 16$	$h = 24$	$h = 40$
Reduced-form innovation to TFP	1.000	0.976	0.783	0.502	0.632	0.537
Agnostic approach	0.519	0.559	0.562	0.502	0.447	0.373

Note: The Table reports the median fraction (across 2000 bootstrap replications) of the h-step ahead forecast error variance of TFP due to surprise technology shocks identified as the reduced-form innovations to TFP and using our agnostic approach.

Table 3.2 shows the contribution of news shocks to the h-step ahead forecast error variance of the series used in estimation. The table also reports the results implied by Barsky & Sims' methodology. Our identified news shocks explain less than 3 percent of the conditional variance of TFP at the one-year horizon and less than 25 percent at the ten-year horizon. They account for more than 35 percent of the forecast error variance of consumption but less than 2 percent of the forecast error variance of output at the one-year horizon. The contribution of news shocks to output variability rises steadily with the forecasting horizon, reaching 38 percent at the ten-year horizon. For hours worked, inflation, stock prices and consumer confidence, the share of the forecast error variance attributed to news shocks never exceeds 16 percent at any given horizon. Compared with the results based on Barsky & Sims' approach, we generally find a smaller contribution of the news shock to aggregate fluctuations at business-cycle frequencies.

Historical decomposition In order to further investigate the importance of news shocks in accounting for business-cycle fluctuations, we simulate the time paths of consumption, output, and hours worked from the estimated VAR assuming that the news shocks are the only stochastic disturbances driving the data. The median results (across

[18]The results shown in the table are the median fractions across the 2000 bootstrap replication.

Table 3.2: Share of Forecast Error Variance attributed to News Shocks. Sample: 1960Q1–2007Q3.

	Horizon					
	$h=1$	$h=4$	$h=8$	$h=16$	$h=24$	$h=40$
	Barsky & Sims' Approach					
TFP	0.000	0.068	0.157	0.311	0.394	0.459
Consumption	0.087	0.207	0.366	0.491	0.503	0.478
Output	0.079	0.073	0.199	0.385	0.428	0.425
Hours	0.419	0.171	0.128	0.155	0.162	0.160
Inflation	0.106	0.172	0.198	0.180	0.175	0.170
Stock Prices	0.040	0.068	0.083	0.112	0.124	0.133
Confidence	0.210	0.223	0.234	0.230	0.218	0.210
	Agnostic Approach					
TFP	0.000	0.027	0.049	0.105	0.154	0.233
Consumption	0.355	0.434	0.419	0.363	0.360	0.381
Output	0.019	0.143	0.219	0.244	0.268	0.313
Hours	0.071	0.091	0.125	0.132	0.137	0.150
Inflation	0.078	0.081	0.072	0.073	0.082	0.096
Stock Prices	0.091	0.100	0.100	0.104	0.112	0.136
Confidence	0.150	0.161	0.148	0.139	0.143	0.151

Note: The table reports the median fraction (across 2000 bootstrap replications) of the h-step ahead forecast error variance of each variable due to news shocks identified using Barsky & Sims' approach (top panel) and our agnostic approach (bottom panel).

2000 bootstrap replications) are depicted in Figure 3.7, where the series are expressed in growth rates. The correlation between the actual and simulated series is high for consumption but fairly low for output and hours worked. News shocks appear to have played a very limited role in explaining post-war U.S. recessions, especially the 1969–1970, 1981–1982, and 1991 recession.

Using the simulated series, we also compute the cross-correlations of the growth rates of consumption, output, and hours. The medians across the 2000 bootstrap replications are reported in Table 3.3. While there is positive comovement between consumption and hours worked in the data, the news shocks identified using Barsky & Sims' methodology imply negative comovement, consistently with the impulse responses shown in the left panels of Figure 3.6. Our agnostic strategy, on the other hand, implies a positive but a much smaller correlation between consumption and hours worked than in the data.

Together with the variance decomposition results discussed above, these observations lead us to conclude that news shocks are unlikely to have been a major driver of business-cycle fluctuations during the period 1960–2007. While this conclusion corroborates that reached by Barsky & Sims (2011), our argument for making such a claim differs from theirs. Indeed, Barsky & Sims (2011) base their conclusion on the fact that consumption co-moves negatively with output and hours worked in response to a news shock, a result that, as we have shown, is largely driven by measurement errors in TFP, just as the

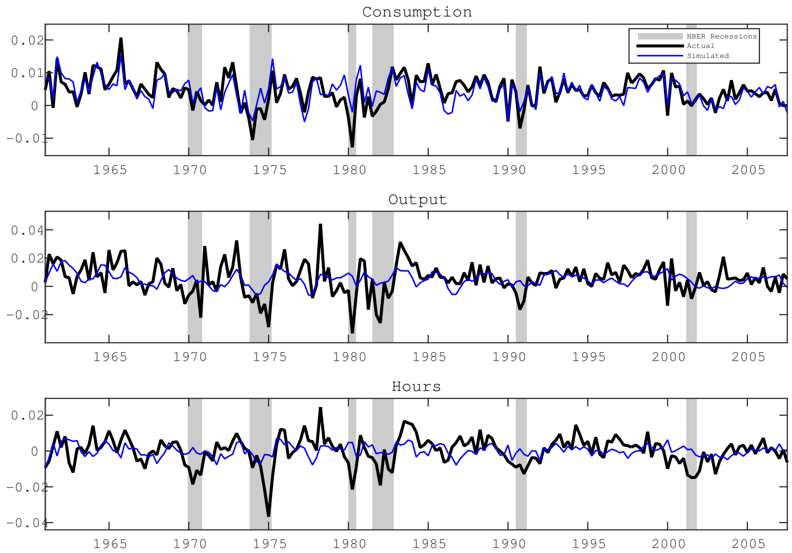

Figure 3.7: Historical decomposition. Sample: 1960Q1-2007Q3.

Notes: The figure shows the actual series (thick black lines) and the ones simulated from the VAR assuming that news shocks are the only stochastic disturbances (thin blue lines). The simulated series are the median across 2000 bootstrap replications. The shaded areas indicate the dates of the U.S. recessions identified by the NBER.

disinflationary effect of the shock. Instead, our conclusion is founded on the fact that news shocks explain only a modest fraction of the variability of output and hours worked at business-cycle frequencies.

Table 3.3: Comovement in the Data and Conditional on News Shocks. Sample: 1960Q1–2007Q3.

	U.S. Data	Barsky & Sims' Approach	Agnostic Approach
$Corr(\Delta \ln C_t, \, \Delta \ln Y_t)$	0.505	0.262	0.456
$Corr(\Delta \ln C_t, \, \Delta \ln N_t)$	0.387	-0.079	0.092
$Corr(\Delta \ln Y_t, \, \Delta \ln N_t)$	0.688	0.853	0.820

Notes: The table reports the historical correlations computed from the data and the ones based on the simulated series (medians across 2000 bootstrap replications) under the assumption that news shocks are the only stochastic disturbances. The variables C_t, Y_t, and N_t denote, respectively, consumption, output, and hours worked.

3.4.2 Sample period 1960Q1–2016Q4

Impulse responses The impulse responses based on the extended sample are reported in Figures 3.8 and 3.9 for the surprise and the news shock, respectively. As before, the left column of each figure shows the results based on Barsky & Sims' methodology while the right column shows the results based on our agnostic approach.

Starting with the surprise technology shock, the results based on the updated sample are very similar to those depicted in the right column of Figure 3.5. The shock has a long-lasting effect on TFP, consumption, and output. Hours worked fall significantly during the year following the shock, but their response is now statistically insignificant during the subsequent horizons. The inflation response to the surprise technology shock is negative by construction during the first eight quarters, and is virtually nil afterward. Stock prices exhibit a positive delayed response, while consumer confidence rises significantly for about ten quarters before returning to its pre-shock level.

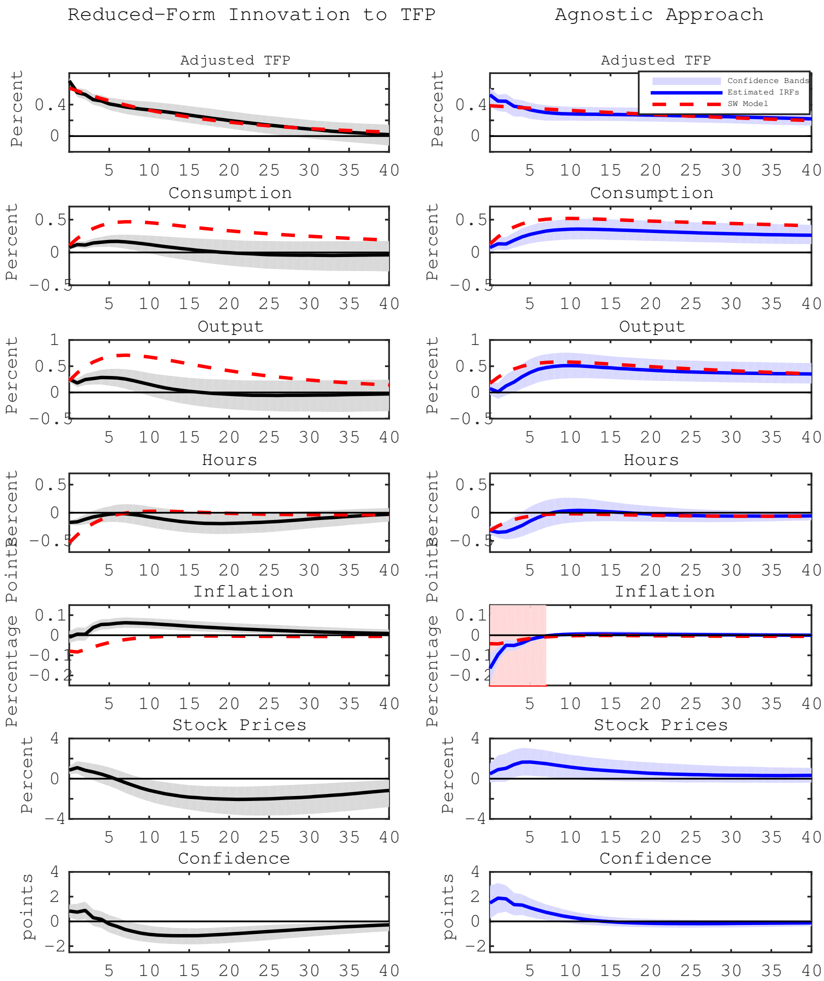

Figure 3.8: Impulse responses to a surprise technology shock. Sample: 1960Q1–2016Q4.

Notes: The figure shows the impulse responses to a surprise technology shock. The solid lines are the median impulse responses estimated based on the reduced-form innovation to TFP (left panels) and on the agnostic approach (right panels). The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications. The shaded red area indicates the horizons at which the inflation response is constrained to be negative. The dashed lines are the impulse responses obtained from the Smets & Wouters (2007) model.

Turning to the responses to the news shock, the left column of Figure 3.9 shows that one of Barsky & Sims' main results, namely the contractionary effect of an anticipated technology shock on output and hours, disappears when we apply their identification strategy to the updated sample. Output increases significantly and persistently but with a delay of three quarters, whereas the response of hours worked is mostly statistically insignificant.[19] The rest of the responses are consistent with those based on the shorter sample. In particular, inflation falls significantly and persistently in response to a good news about future technology.

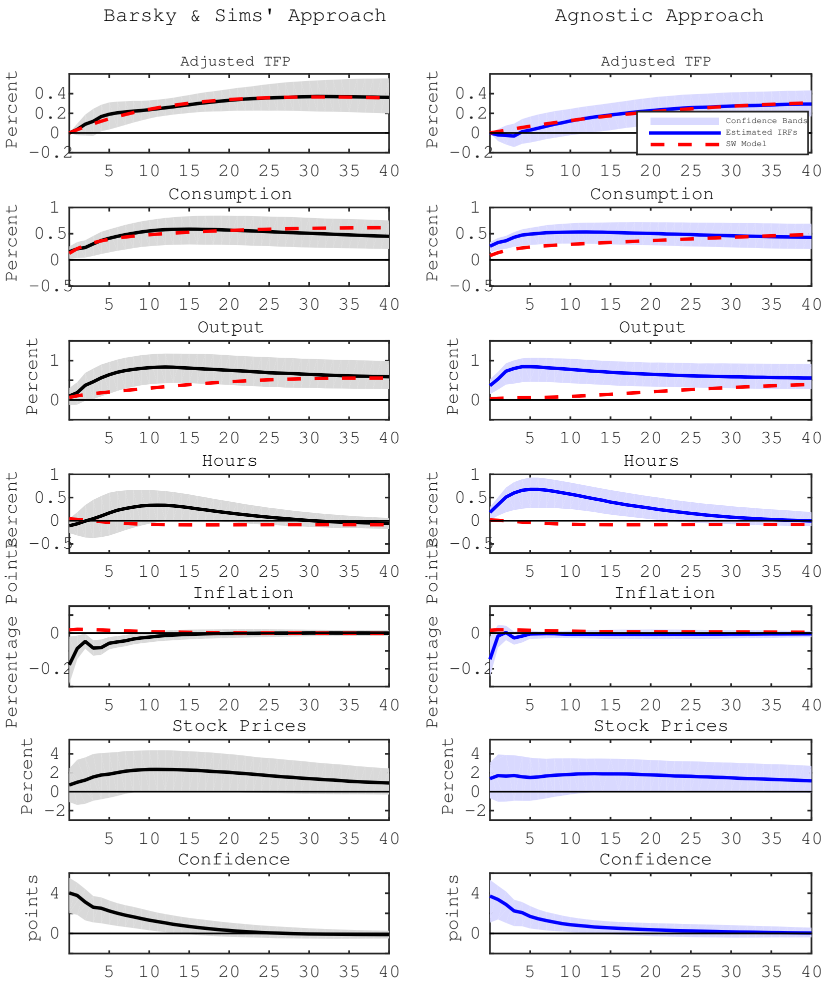

Figure 3.9: Impulse responses to a news shock. Sample: 1960Q1–2016Q4.

Notes: The figure shows the impulse responses to a news shock. The solid lines are the median impulse responses estimated based on Barsky and Sims' approach (left panels) and on the agnostic approach (right panels). The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications. The dashed lines are the impulse responses obtained from the Smets & Wouters (2007) model.

The results based on our agnostic identification strategy show important differences both with respect to those implied by Barsky & Sims' methodology and those based on the shorter sample. First, TFP exhibits a much more inertial response to the shock, starting

[19]Kurmann & Sims (2017) attribute the difference in results to the revisions in Fernald (2014)'s adjusted TFP series, and interpret the lack of robustness as an indication of the presence of measurement errors. However, we find Barsky and Sims' original results to remain largely unchanged when we apply their approach to the 2016 vintage of TFP but using the same sample period as in their paper (i.e., 1960Q1–2007Q3). This suggests that the revisions in TFP are unlikely to be driving the difference in results obtained using the 2007 and 2016 vintages of TFP.

to increase in a statistically significant manner only after about three years. This slowly diffusing process contrasts with the rapid increase in TFP estimated based on the shorter sample and using Barsky & Sims' methodology. Second, consumption, output, and hours worked increase significantly and persistently in response to the news shock. This simultaneous increase in macroeconomic aggregates — indicative of positive comovement — occurs well before TFP starts to rise; a result that corroborates Beaudry & Portier (2006)'s original findings. Third, inflation falls in response to the shock but its response exhibits very little persistence and is (barely) statistically significant only on impact. In other words, the disinflation puzzle appears to be much less acute under our identification strategy. Finally, unlike the results based on the shorter sample, the estimated impulse responses match rather poorly those implied by the Smets & Wouters (2007) model.

Variance decomposition Variance decomposition results for the updated sample are reported in Table 3.4. One of the striking differences with the results based on the shorter sample and on Barsky & Sims' methodology is that news shocks account for a relatively large fraction of the forecast error variance of output and hours worked at short horizons. At the one-year horizon, this fraction amounts to 42 percent for output and 28 percent for hours. At business-cycle frequencies, the contribution of news shocks to the variability of consumption, output, and hours worked ranges roughly between 40 and 60 percent. On the other hand, news shocks continue to explain a small fraction of the forecast error variance of inflation, stock prices, and, consumer confidence at business-cycle frequencies. Our agnostic approach continues to attribute a smaller role to news shocks in accounting for the conditional variance of these variables than does Barsky & Sims' methodology.

Historical decomposition Figure 3.10 shows the actual growth rates of consumption, output, and hours worked, along with their counterparts based on the artificial series simulated under the assumption that news shocks are the only underlying disturbances. The actual and simulated series for output and hours worked are more highly correlated than in the shorter sample, while actual and simulated consumption growth continue to track each other very closely. The figure also shows that news shocks account for a significant share of the decline in consumption, output, and hours worked during all of the U.S. recessions after 1973, including the Great Recession.

Table 3.5 reports the median cross-correlations of the growth rates of consumption, output, and hours worked based on the simulated series. The table confirms that the negative comovement between consumption and hours worked documented by Barsky & Sims (2011) vanishes when their methodology is applied to the extended sample period. The implied correlation, however, is lower than in the data. A much stronger positive comovement between consumption, output, and hours worked is obtained conditional on the news shocks identified using our agnostic strategy. These findings, along with the variance-decomposition results, suggest that news shocks have become an important

Table 3.4: Share of Forecast Error Variance attributed to News Shocks. Sample: 1960Q1–2016Q4.

	Horizon					
	$h = 1$	$h = 4$	$h = 8$	$h = 16$	$h = 24$	$h = 40$
	Barsky & Sims' Approach					
TFP	0.000	0.024	0.087	0.200	0.315	0.455
Consumption	0.137	0.226	0.355	0.493	0.531	0.527
Output	0.041	0.125	0.257	0.427	0.473	0.482
Hours	0.067	0.070	0.101	0.153	0.162	0.161
Inflation	0.188	0.206	0.219	0.212	0.193	0.181
Stock Prices	0.059	0.085	0.119	0.166	0.186	0.320
Confidence	0.291	0.360	0.388	0.372	0.342	0.210
	Agnostic Approach					
TFP	0.000	0.012	0.025	0.073	0.144	0.242
Consumption	0.372	0.498	0.534	0.597	0.484	0.452
Output	0.203	0.416	0.494	0.486	0.461	0.438
Hours	0.119	0.280	0.367	0.406	0.390	0.371
Inflation	0.130	0.103	0.092	0.088	0.096	0.105
Stock Prices	0.102	0.113	0.124	0.142	0.152	0.171
Confidence	0.251	0.296	0.288	0.265	0.252	0.246

Note: The table reports the median fraction (across 2000 bootstrap replications) of the h-step ahead forecast error variance of each variable due to news shocks identified using Barsky & Sims' approach (top panel) and our agnostic approach (bottom panel).

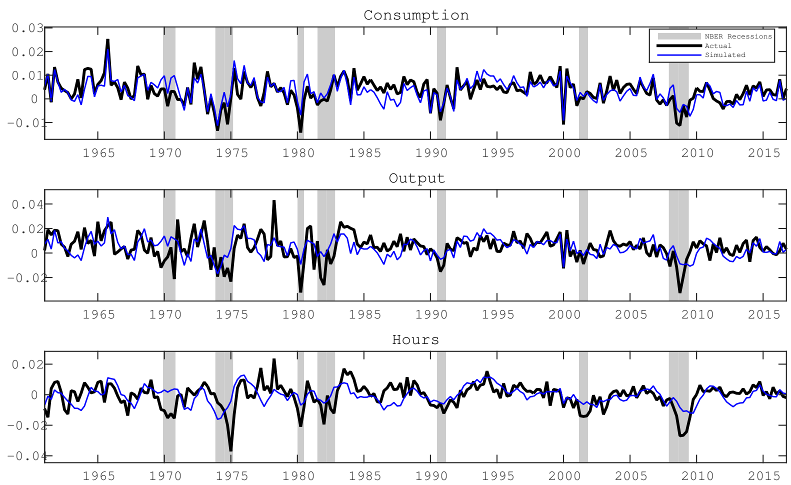

Figure 3.10: Historical decomposition. Sample: 1960Q1-2016Q4.

Notes: The figure shows the actual series (thick black lines) and the ones simulated from the VAR assuming that news shocks are the only stochastic disturbances (thin blue lines). The simulated series are the median across 2000 bootstrap replications. The shaded areas indicate the dates of the U.S. recessions identified by the NBER.

driver of business-cycle fluctuations in recent years. In this respect, our agnostic identification strategy provides a sharply contrasting conclusion to that based on Barsky & Sims' methodology or variants of it used in recent empirical studies.

Table 3.5: Comovement in the Data and Conditional on News Shocks. Sample: 1960Q1–2016Q4.

	U.S. Data	Barsky & Sims' Approach	Agnostic Approach
$Corr(\Delta \ln C_t,\ \Delta \ln Y_t)$	0.538	0.771	0.894
$Corr(\Delta \ln C_t,\ \Delta \ln N_t)$	0.391	0.298	0.730
$Corr(\Delta \ln Y_t,\ \Delta \ln N_t)$	0.667	0.731	0.878

Notes: The table reports the historical correlations computed from the data and the ones based on the simulated series (medians across 2000 bootstrap replications) under the assumption that news shocks are the only stochastic disturbances. The variables C_t, Y_t, and N_t denote, respectively, consumption, output, and hours worked.

3.5 Robustness: Systematic Measurement Errors

The identification strategy proposed in this paper relies on the commonly used assumption that news shocks do not affect measured TFP contemporaneously. However, to the extent that non-technological shocks affecting TFP subsume systematic measurement errors in factor utilization, the zero-impact assumption may become unwarranted, since news shocks could affect measured TFP through their effects on input utilization. Based on the latter argument, Kurmann & Sims (2017) relax the assumption that measured TFP does not react contemporaneously to news shocks and identify these shocks solely based on the Max Share criterion described above. Using this strategy, Kurmann & Sims (2017) find very similar effects of the news shock to those reported by Barsky & Sims (2011). In particular, they find that consumption rises while hours worked and inflation decline in response to a favorable news shock. Importantly, they show that these results remain robust when they update the sample to include the 2016 vintage of Fernald's adjusted TFP series.

A crucial assumption of Kurmann & Sims' identification scheme is that the news shock is not orthogonalized with respect to the surprise technology shock. Because the latter is typically identified as the reduced-form innovation to TFP, imposing orthogonality with respect to this shock necessarily implies that the contemporaneous response of TFP to the news shock is nil,[20] which is precisely the restriction that Kurmann & Sims (2017) aim to relax (and to which we henceforth refer as the "zero-impact" restriction). This in turn suggests that Kurmann & Sims' strategy is likely to confound surprise and anticipated technological shocks, as both shocks affect TFP in the short and in the long run, making it impossible — without further assumptions — to disentangle their respective contribution to the forecast error variance of TFP at any given horizon.

Our agnostic approach, on the other hand, allows us to relax the zero-impact restriction while still imposing the orthogonality of the news shock with respect to the surprise shock, since the latter is identified via sign restrictions. The prior identification of the

[20] Assuming again that TFP is ordered first in y_t, the impulse vector associated with the surprise technology shock has zeros everywhere except for the first element. For this impulse vector to be orthogonal to the one associated with the news shock, the latter must have zero as its first element.

surprise shock enables us to identify the news shock by maximizing its contribution to the remainder of the forecast error variance of TFP at any given (range of) horizon(s).

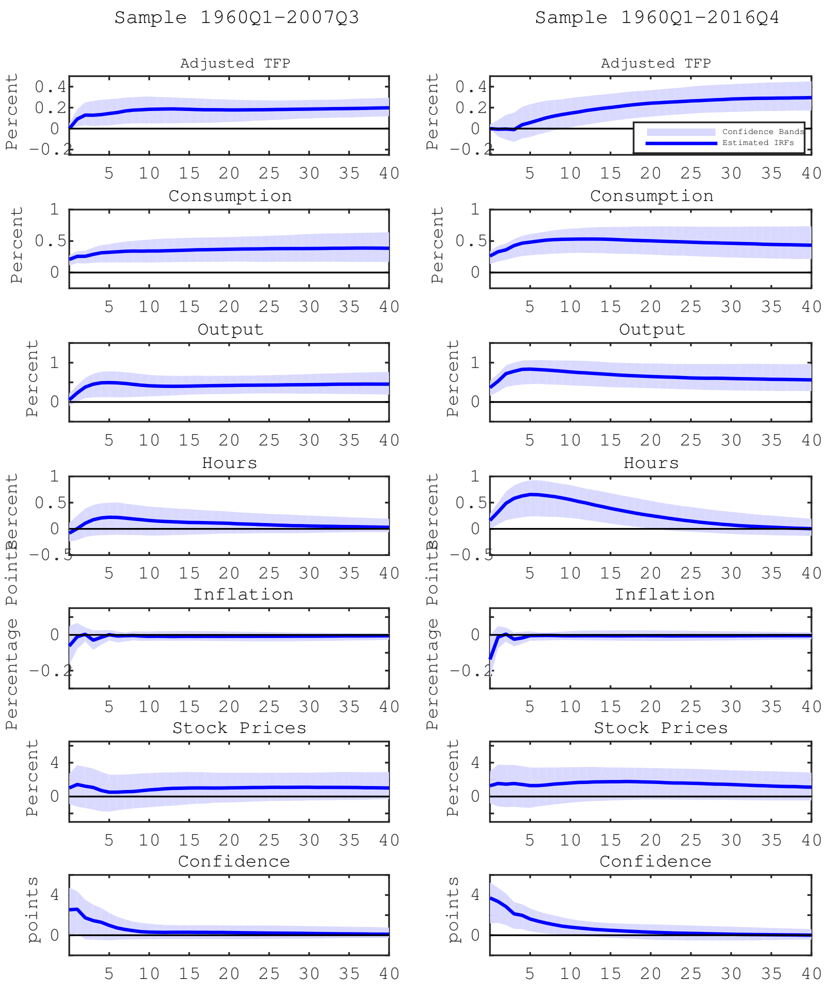

Figure 3.11: Impulse responses to a news shock: Relaxing the zero-impact restriction.

Notes: The figure shows the impulse responses to a news shock estimated using the agnostic strategy under the assumption that the news shock can affect TFP on impact. The solid lines are the median impulse responses. The 68 percent confidence bands are the bias-corrected bootstrap confidence intervals computed using Kilian (1998)'s procedure with 2000 replications.

We apply this variant of our agnostic approach to the two sample periods considered in the previous section. To do so, we relax the restriction $\gamma_2(1) = 0$ in the maximization problem described in Section 3.3.2. The impulse responses to a news shock based on this approach are shown in Figure 3.11. Interestingly, even though the zero-impact restriction is relaxed, the median initial response of adjusted TFP turns out to be equal to zero — with very little sampling uncertainty — regardless of the sample period. The estimated response of TFP during the subsequent quarters is remarkably similar to that estimated under the zero-impact restriction. This can be seen by comparing the upper left panel of Figure 3.11 with the upper right panel of Figure 3.6 for the 1960Q1–2007Q3 sample period, and the upper right panels of Figures 3.11 and 3.9 for the 1960Q1–2016Q4 sample period. Not surprisingly (given this similarity), the impulse responses of the remaining variables are hardly affected when the zero-impact restriction is relaxed. In particular, hours worked continue to be unresponsive to the news shock in the sample ending in 2007, and to increase significantly and persistently along with consumption and output in the updated sample, while the disinflation puzzle essentially vanishes in both samples. These findings contradict those reported by Kurmann & Sims (2017) and suggest that their identified news shock is partly picking up the effects of the unanticipated technology shock.

We also find that the variance-decomposition results and the historical decomposition of macroeconomic aggregates exhibit very little sensitivity to the zero-impact restriction,[21] thus confirming our main conclusions: news shocks contributed very little to business-cycle fluctuations during the 1960Q1–2007Q3 period, but their importance has increased significantly in recent years.

3.6 Conclusion

Much of the recent VAR-based evidence on the effects of news shocks about future productivity casts doubt on the plausibility and importance of TFP-news-driven business cycles, as these shocks are found to generate negative comovement between consumption and hours worked. Another robust finding of this literature is that favorable news shocks tend to be associated with sharp and persistent declines in inflation.

In this paper, we have shown that these conclusions are spurious and are largely due to the presence of measurement errors in TFP. We have documented the severity of these errors by examining the effects of unanticipated technology shocks, usually identified as the reduced-form innovations to TFP. We found these effects to be inconsistent with the interpretation of unanticipated technological disturbances as supply shocks. We have then proposed an agnostic identification strategy that is robust to measurement errors, successfully isolating the technological component of TFP. We found no evidence of negative comovement between consumption and hours worked conditional on a news shock, and the disinflation puzzle essentially disappears under our identification strategy. Importantly, we found that news shocks have become a major source of business-cycle fluctuations in recent years, consistently with Beaudry & Portier (2006)'s original view.

News about TFP, however, are clearly not the only factor that can cause changes in agents' expectations. Some recent studies have empirically examined the importance of changes in expectations caused by factors unrelated to TFP, such as news about investment-specific technology (e.g., Ben Zeev & Khan (2015)) or sentiments (e.g., Beaudry et al. (2011), Levchenko & Pandalai-Nayar (2015) and Fève & Guay (2016)). The identification of these shocks, however, usually relies on the prior identification of TFP news shocks, which implies that the empirical approaches developed in this strand of the literature are

[21] To conserve space, these results are not reported but are available upon request.

also likely to be plagued by measurement errors in TFP. By correctly identifying TFP news shocks, the empirical strategy developed in this paper can therefore help shed light on the relative importance of non-TFP news shocks for aggregate fluctuations.

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Appendices

A Appendix to Chapter 1

A.1 Appendix to the Theoretical Model

Steady-State of the Model

We solve the model using third-order perturbation in the neighborhood of the steady-state. Solving for the deterministic steady-state is therefore necessary. The model equations in steady-state are given by the following equations where variables without a time subscript denote steady-state values and where I abstract from including equations of the time-varying standard deviations and policy instruments whose steady-state values are naturally their unconditional means, and the government spending-to-GDP ratio is set to 0.1984 following Fernández-Villaverde et al. (2011):

$$V = c^{\nu} (1-n)^{1-\nu} \tag{9}$$

$$\frac{\nu(1-\gamma)}{\theta} c^{\frac{\nu(1-\gamma)}{\theta}-1} (1-n)^{\frac{(1-\nu)(1-\gamma)}{\theta}} = \lambda(1+\tau^c) \tag{10}$$

$$\frac{1-\nu}{\nu} \frac{(1+\tau^c)c}{(1-\tau^n)(1-n)} = w \tag{11}$$

$$y = ak^{\alpha} n^{1-\alpha} \tag{12}$$

$$mc = \frac{\sigma-1}{\sigma} \tag{13}$$

$$c + i + g = y \tag{14}$$

$$a = 1 \tag{15}$$

$$\left[\delta_0 + \delta_1(u-1) + \frac{\delta_2}{2}(u-1)^2\right] K = i \tag{16}$$

$$\delta_0 k^b = i \tag{17}$$

$$r_k(1-\tau^k) = q\left[\delta_1 + \delta_2(u-1)\right] \tag{18}$$

$$q\left\{1 - \beta\left[1 - \delta_0 - \delta_1(u-1) - \frac{\delta_2}{2}(u-1)^2\right]\right\} = \beta r_k(1-\tau^k)u \tag{19}$$

$$q^b = \beta\left[\delta_0 \tau^k + (1-\delta_0)q^b\right] \tag{20}$$

$$q^b + q = 1 \tag{21}$$

$$mc = \left(\frac{r_k}{\alpha}\right)^{\alpha} \left(\frac{w}{1-\alpha}\right)^{1-\alpha} \tag{22}$$

$$(1-\alpha) r_k k = \alpha w n \tag{23}$$

$$k = uK \tag{24}$$

$$Q^l = \left(\frac{\beta}{\pi}\right)^l, \ l = 0, \cdots, L \tag{25}$$

I choose the value of δ_1 such that the capital utilization rate is equal to one in the long-run, by combining equations (18) and (19), and imposing $u = 1$. We obtain $\delta_1 = \frac{1}{\beta} - (1 - \delta_0)$. Monetary policy authorities have an annual inflation target of 2% and therefore set π equal to $1.02^{\frac{1}{4}}$, which permits to get steady-state values of $Q^l, \ l = 0, \cdots, L$ based on (25). We choose a value of ν that ensures $n = \frac{1}{3}$. Finally, the government spending-to-GDP ratio is set to 0.1984 as stated earlier.

Given this setup, the steady-state values of remaining variables can easily be computed. First of all, use (20) to get q^b and substitute it into (21) to get q. Then use q and (18) to compute r_k. Since mc is entirely determined by σ through (13), we can obtain w by substituting mc and r_k into (22). Once we have w, (23) allows to compute k, and (12) combined with (15) yields y. From (24) and $u = 1$ we have that $K = k$, which allows us to get i through (16). With i, we have k^b using (17). Using (14), i, y and $g = 0.1984y$, we can compute c. (11) permits to get the value of ν which ensures that in the steady-state, households use one-third of their time at work ($n = \frac{1}{3}$) as assumed above. Finally, (10) and (9) can be used to obtain λ and V.

Supplemental Figures

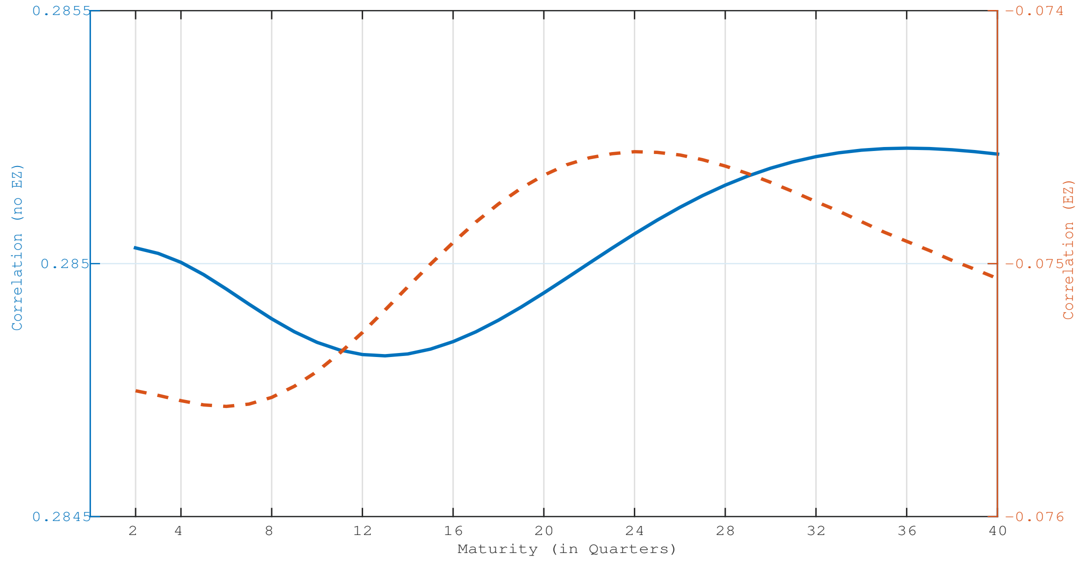

Figure A.1: Correlation between current consumption and expected holding-period return as a function of maturity.

Notes: Right axis (green dotted line): model featuring recursive preferences. Left axis (solid blue line): model featuring standard preferences.

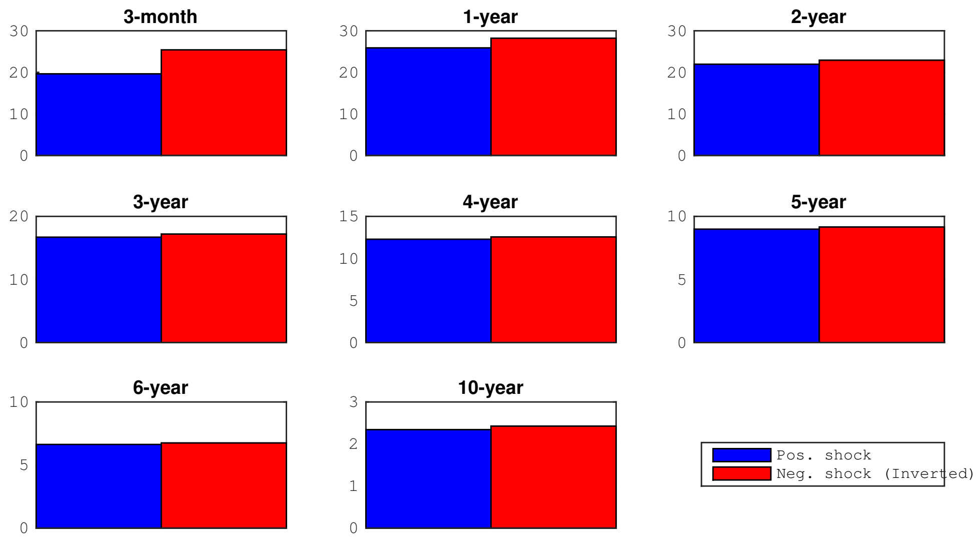

Figure A.2: Impact effect of a two-standard deviation policy risk shock (in annualized basis points) for different maturities.

Notes: Blue bars represent effects of a positive shock and red bars—which mostly represent negative values but are inverted for the sake of comparison—represent effects of negative shocks.

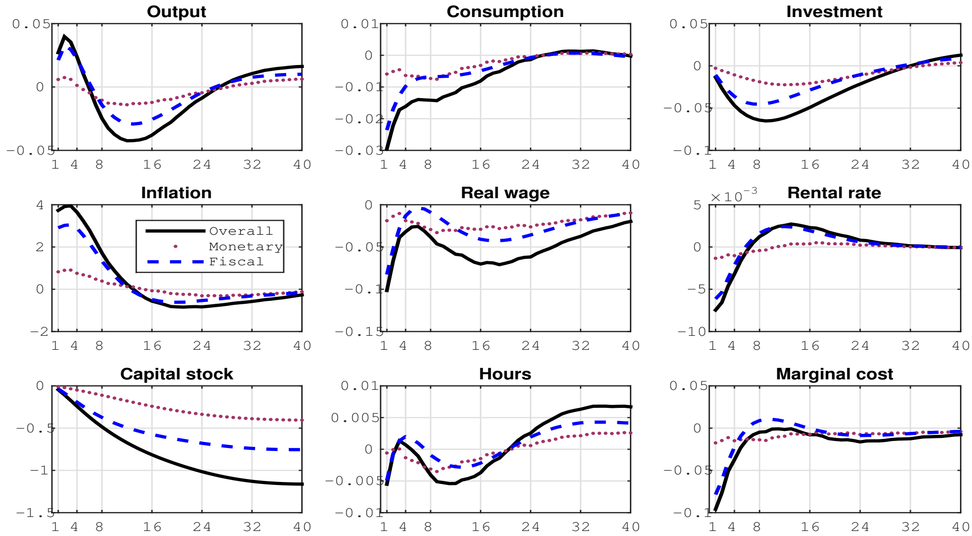

Figure A.3: RBC effects of a two-standard deviation policy risk shock in the model with investment adjustment costs.

Notes: All responses are in percent except for those of inflation that are in annualized basis points. The solid black lines depict the effects of the overall policy risk shock; the purple dotted lines depict the effects of the monetary policy risk shock, and the blue dashed line represent the effect of the fiscal policy risk shock. Horizontal axes represent quarters.

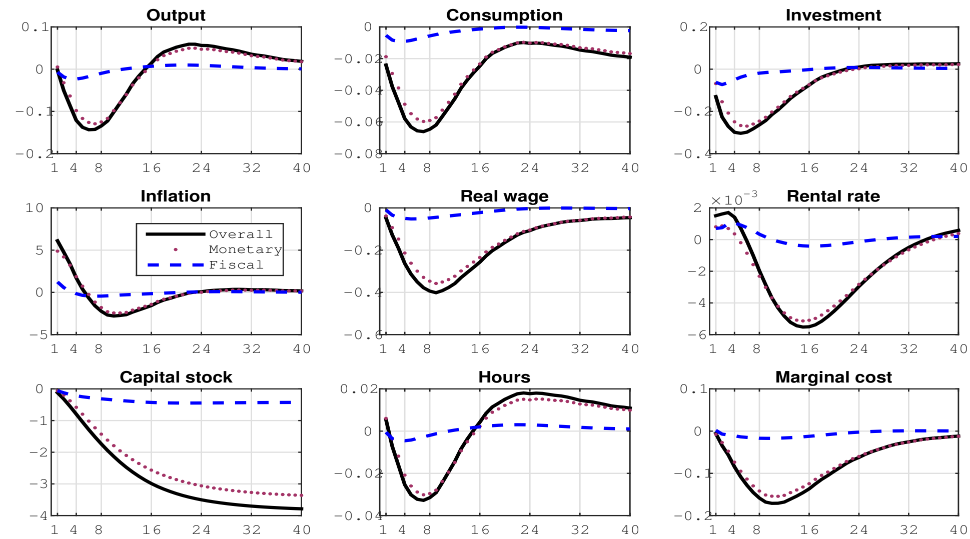

Figure A.4: RBC effects of a two-standard deviation policy risk shock: no investment adjustment costs.

Notes: All responses are in percent except for those of inflation that are in annualized basis points. The solid black lines depict the effects of the overall policy risk shock; the purple dotted lines depict the effects of the monetary policy risk shock, and the blue dashed line represent the effect of the fiscal policy risk shock. Horizontal axes represent quarters.

A.2 Appendix to the Empirical Model

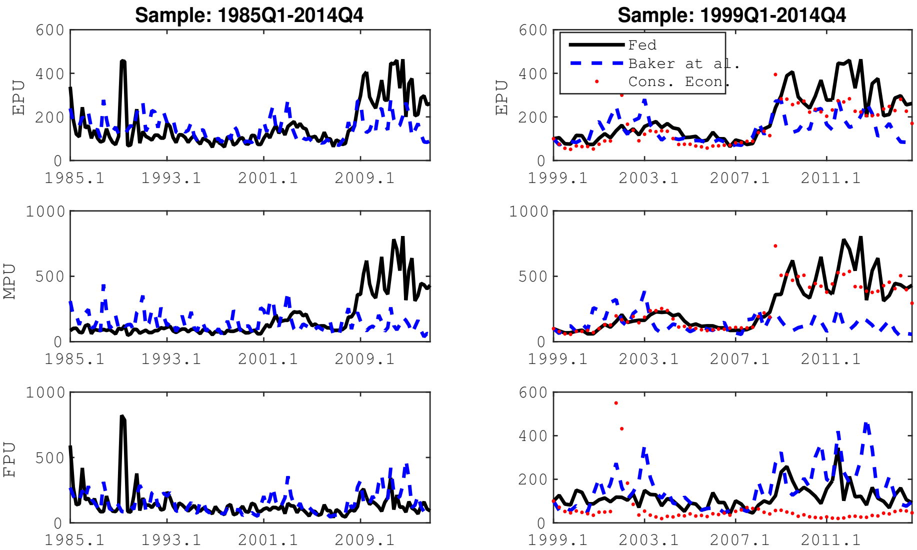

Figure A.5: Three measures of policy uncertainty.

Notes: Based on the Philadelphia Fed's data (solid black lines), on Consensus Economics' (dotted red lines) and Baker et al. (2013)'s measure (dashed blue lines). EPU, MPU, and FPU stand for economic, monetary, and fiscal policy uncertainty respectively.

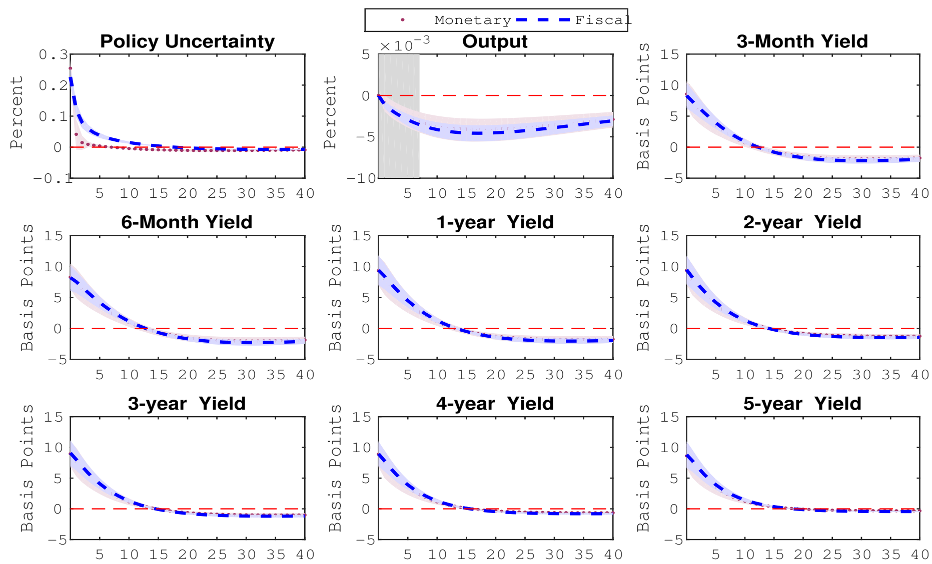

Figure A.6: Impulse responses to one-standard deviation policy uncertainty shocks: Baker et al. (2013)'s data (1985Q1-2014Q4).

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

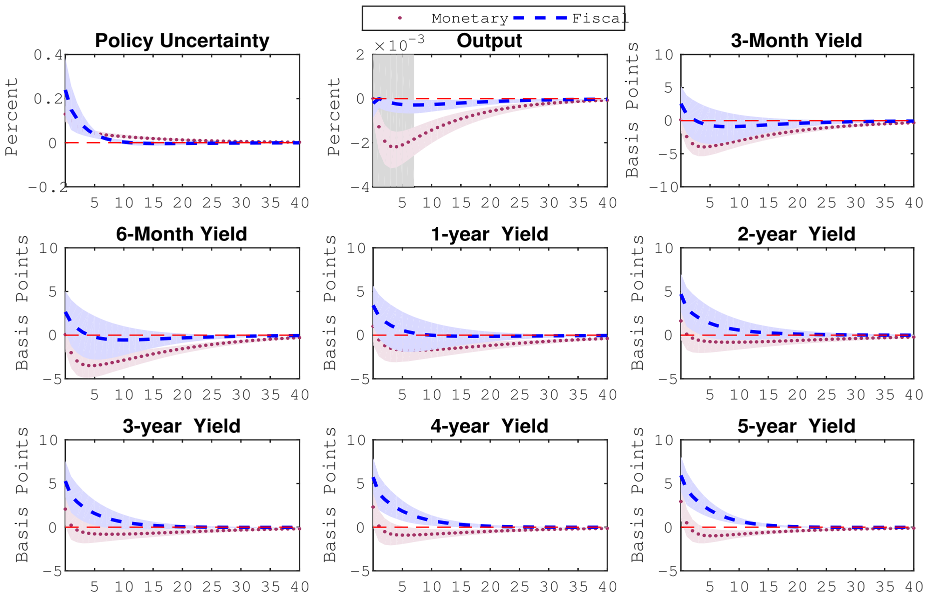

Figure A.7: Impulse responses to one-standard deviation policy uncertainty shocks: Consensus Economics' data (1999Q1-2014Q4).

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

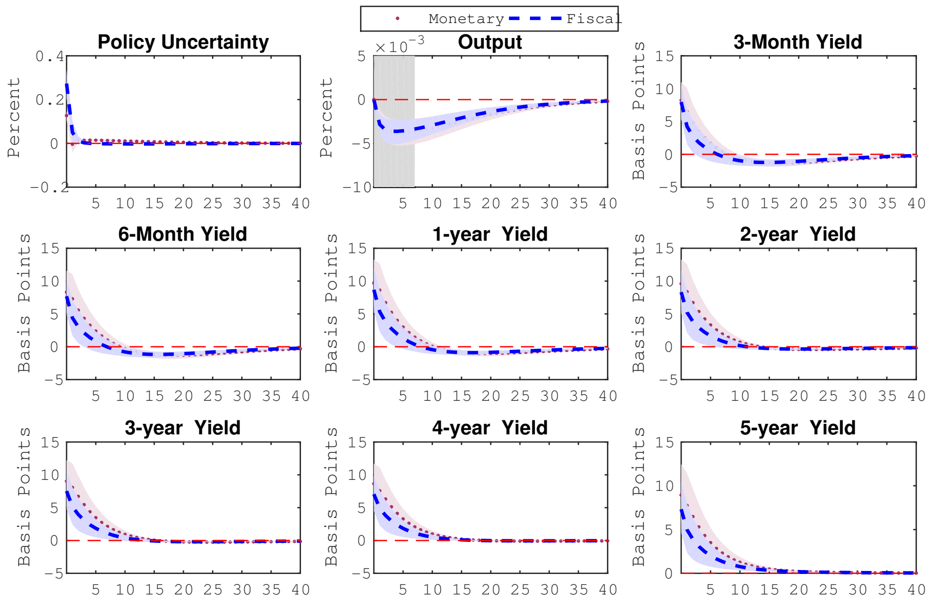

Figure A.8: Impulse responses to one-standard deviation policy uncertainty shocks: Benchmark system, sub-sample 1981Q3-1998Q4.

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

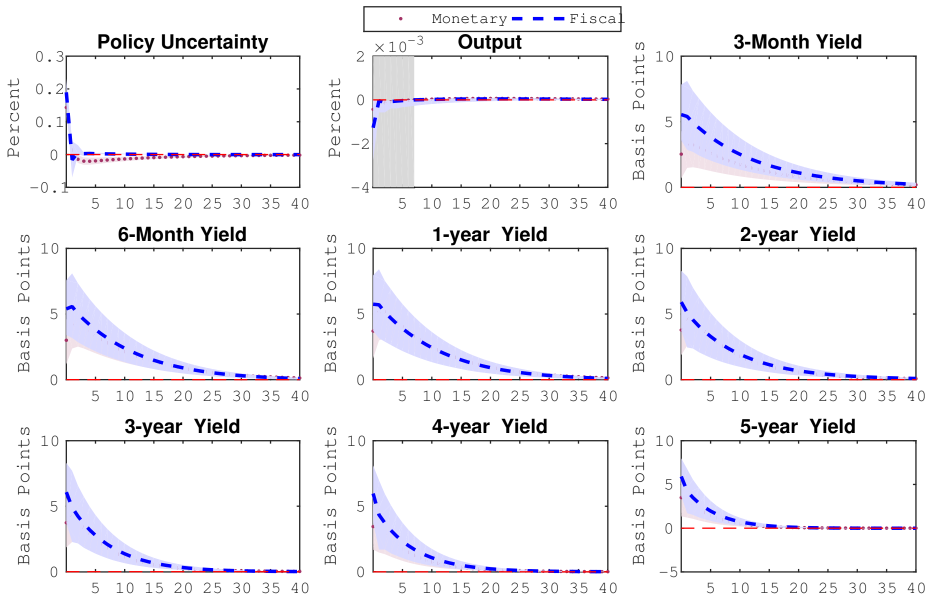

Figure A.9: Impulse responses to one-standard deviation policy uncertainty shocks: Benchmark system, sub-sample 1999Q1-2014Q4.

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

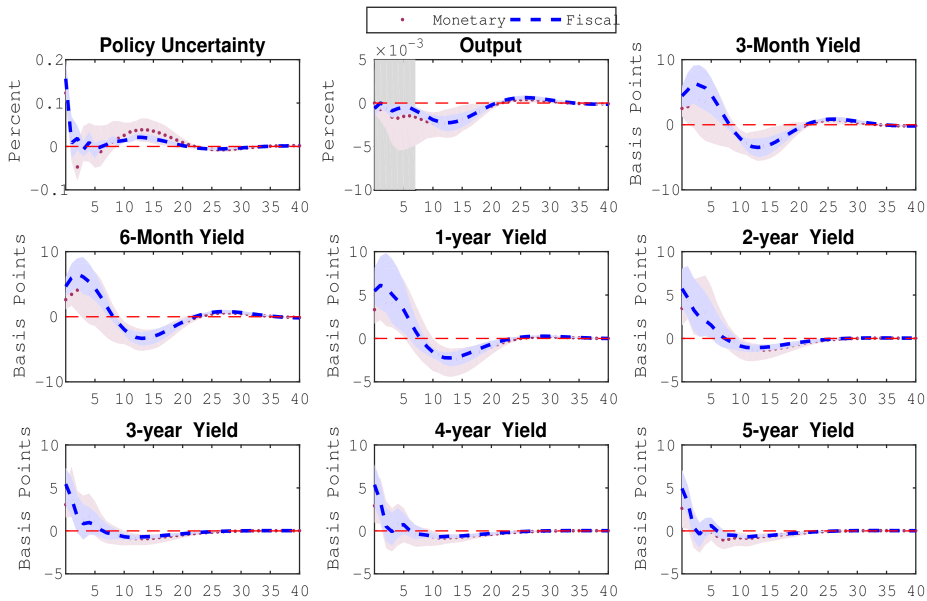

Figure A.10: Impulse responses to one-standard deviation policy uncertainty shocks: Benchmark system with three lags in the reduced-form VAR.

Notes: The purple dotted lines depict the effects of the monetary policy uncertainty shock, and the blue dashed line represent the effect of the fiscal policy uncertainty shock. Horizontal axes represent quarters.

B Appendix to Chapter 2

B.1 Forward Risk Premium Differentials in a simple Two-period Portfolio Choice Model

Here I present the derivations of the portfolio choice model presented in Section 2.2.2. Consider two governments, "domestic" and "foreign", who issue two types of securities each, all denominated in the home currency. The first two securities are one-period maturity government bonds that pay interest r_t when issued by the domestic country (and $r_t^\star$ when issued by the foreign country). The two other securities are forward contracts which, if held by an investor in the first period, promise to pay interest f_t (respectively $f_t^\star$ if it is issued by the foreign government) on bonds issued in the second period. Consider a domestic risk-averse portfolio manager choosing among these four securities in the first period and two securities in the second. Each period, he faces costs l that reduce his wealth and that could be attributed to poor economic and institutional characteristics of the domestic country. The losses due to poor economic characteristics derive from low liquidity of the local financial market, high reinvestment risk, high inflation, high taxes triggered by bad fiscal performance, (and exchange rate fluctuations if the assets are denominated in different currencies as it is the case in the empirical analysis) etc. These costs are proportional to the amount of domestic transactions carried out by the investor. Securities issued by the foreign government are considered as benchmark in the market and the costs related to them are normalized to zero.

The portfolio manager maximizes an investor's utility function that depends positively on the expected wealth, $E_t[w_{t+2}]$ and negatively on its variance, $Var_t[w_{t+2}]$:

$$Max\ U\{E_t[w_{t+2}], Var_t[w_{t+2}]\},\ U_1 > 0,\ U_2 < 0. \tag{26}$$

The portfolio manager takes all investment decisions in the first period, even decisions regarding the composition of the portfolio of bonds in the second period. Let's assume that the decision on the portfolio composition in period two is based on the expected values of all information needed by the portfolio manager (for instance expected wealth, rates, costs and probabilities of default). In the first period, the portfolio manager therefore chooses the fractions θ_t^{ds} and θ_t^{fs} of his wealth w_t to allocate to domestic and foreign government bonds respectively; the fractions θ_t^{df} and θ_t^{ff} to allocate to domestic and foreign forward contracts respectively; and the fractions θ_{t+1}^{ds} and $1-\theta_{t+1}^{ds}$ of his expected second period's wealth ($E_t[w_{t+1}]$) to allocate to domestic and foreign government bonds respectively. The portfolio manager faces the following budget constraint:

$$\theta_t^{ds} + \theta_t^{df} + \theta_t^{fs} + \theta_t^{ff} = 1 \tag{27}$$

As Bernoth et al. (2012), I assume that domestic securities are subject to the risk of partial default, but foreign assets are risk-free. With a probability of $1 - P(x_t),\ 0 \leq P(x_t) \leq 1$, the domestic government will default on its debt, repaying only a fraction, $\alpha_t \in (0, 1)$, of it. x_t denotes the set of variables that influence this probability. In the following, I will use P_t instead for convenience and denote its period's two expected value by P_{t+1}. The expected wealth and its variance are given by:

$$
\begin{aligned}
E_t[w_{t+2}] = (1+E_t r_{t+1})\,\theta^{ds}_{t+1} E_t[w_{t+1}] P_{t+1} + \alpha_{t+1}\,(1+E_t r_{t+1})\,\theta^{ds}_{t+1} E_t[w_{t+1}]\,(1-P_{t+1}) \\
- \theta^{ds}_{t+1} E_t[w_{t+1}] l_{t+1} + (1+f_t)\,\theta^{df}_t w_t P_{t+1} + \alpha_{t+1}\,(1+f_t)\,\theta^{df}_t w_t\,(1-P_{t+1}) \\
+ \left(1+E_t r^{\star}_{t+1}\right)\left(1-\theta^{ds}_{t+1}\right) E_t[w_{t+1}] + (1+f^{\star}_t)\, w_t \theta^{ff}_t
\end{aligned} \tag{28}
$$

$$
Var_t[w_{t+2}] = (1-\alpha_{t+1})^2 \left[(1+E_t r_{t+1})\,\theta^{ds}_{t+1} E_t[w_{t+1}] + (1+f_t)\,\theta^{df}_t w_t\right]^2 P_{t+1}\,(1-P_{t+1}) \tag{29}
$$

where the expected wealth at the end of the first period, $E_t[w_{t+1}]$, is given by

$$
E_t[w_{t+1}] = (1+r_t)\,\theta^{ds}_t w_t P_t + \alpha_t\,(1+r_t)\,\theta^{ds}_t w_t\,(1-P_t) - \left(\theta^{ds}_t + \theta^{df}_t\right) w_t l_t + (1+r^{\star}_t)\,\theta^{fs}_t w_t \tag{30}
$$

The portfolio manager's problem is, thus, given by (26) subject to (27) [and (28), (29) and (30)].
Let $A_{t+1} = (1+E_t r_{t+1})\,P_{t+1} + \alpha_{t+1}\,(1+E_t r_{t+1})\,(1-P_{t+1}) - l_{t+1} - \left(1+E_t r^{\star}_{t+1}\right)$. Let also $\Phi_t = -2w_t U_2/U_1$ be the coefficient of relative risk aversion for the investor.
The first-order condition with respect to θ^{ds}_{t+1} yields the following equation:

$$
\theta^{ds}_{t+1} = \frac{A_{t+1} w_t}{(1-\alpha_{t+1})^2\,(1+E_t r_{t+1})^2\,P_{t+1}\,(1-P_{t+1})\,E_t[w_{t+1}]\Phi_t} - \frac{(1+f_t)\,w_t}{(1+E_t r_{t+1})\,E_t[w_{t+1}]}\theta^{df}_t \tag{31}
$$

And the first-order condition with respect to θ^{df}_t is given by:

$$
\begin{aligned}
(1+f_t)\,[P_{t+1} + \alpha_{t+1}\,(1-P_{t+1})] - (1+f^{\star}_t) - \left(A_{t+1}\theta^{ds}_{t+1} + 1 + E_{t+1} r^{\star}_{t+1}\right) l_t \\
= (1-\alpha_{t+1})^2 \left[(1+E_{t+1} r_{t+1})\,\theta^{ds}_{t+1} E_{t+1} w_{t+1} + (1+f_t)\,w_t \theta^{df}_t\right] \\
\times \left[(1+f_t) - (1+E_{t+1} r_{t+1})\,\theta^{ds}_{t+1} l_t\right] P_t P_{t+1} \frac{\Phi_t}{w_t}
\end{aligned} \tag{32}
$$

Combining equations (31) and (32), and doing some simple algebra yields:

$$
\left(1+E_{t+1} r^{\star}_{t+1}\right) l_t + \frac{A_{t+1}\,(1+f_t)}{1+E_{t+1} r_{t+1}} = (1+f_t)\,[P_{t+1} + \alpha_{t+1}\,(1-P_{t+1})] - (1+f^{\star}_t) \tag{33}
$$

Substituting A_{t+1} in the previous equation yields:

$$
\frac{1+f^{\star}_t}{1+f_t} - \frac{1+E_t r^{\star}_{t+1}}{1+E_t r_{t+1}} = \frac{l_{t+1}}{1+E_t r_{t+1}} - \frac{1+E_t r^{\star}_{t+1}}{1+f_t} l_t \tag{34}
$$

The last equation can in turn be rearranged to obtain the forward risk premium differential on domestic and foreign government bonds as follows (equation 2.9 in the main text):

$$
(f_t - E_t r_{t+1}) - \left(f^{\star}_t - E_t r^{\star}_{t+1}\right) = l_t + \Psi_t \tag{35}
$$

where $\Psi_t = l_t \left[E_t r^{\star}_{t+1} + \left(1 - E_t r^{\star}_{t+1}\right) E_t r_{t+1}\right] - l_{t+1}\left(1 + f_t\right) + f^{\star}_t E_t r_{t+1} - f_t E_t r^{\star}_{t+1}$ is a term that involves various covariances among the different interest rates as well as the expected cost.

B.2 Results on Convergence Analysis with Alternative Measure of Risk Premiums

Table B.6: Correlations of forward risk premiums across countries and by maturity (using the alternative measure of forward risk premiums).

	AUS	BEL	CAN	FRA	ITA	MLT	NOR	SWE	USA
AUS	1.00								
BEL	0.69	1.00							
CAN	0.64	0.74	1.00						
FRA	0.68	1.00	0.75	1.00					
ITA	0.67	0.97	0.75	0.98	1.00				
MLT	0.66	0.94	0.66	0.93	0.91	1.00			
NOR	0.58	0.88	0.60	0.89	0.86	0.93	1.00		
SWE	0.61	0.87	0.62	0.87	0.85	0.89	0.87	1.00	
USA	0.52	0.61	0.82	0.61	0.63	0.46	0.35	0.27	1.00
CRO	**-0.51**	**-0.62**	**-0.81**	**-0.62**	**-0.63**	**-0.58**	**-0.58**	**-0.50**	**-0.69**
HKG	**0.46**	**0.55**	**0.81**	**0.55**	**0.57**	**0.42**	**0.36**	**0.22**	**0.95**
HUN	**-0.30**	**-0.31**	**-0.56**	**-0.31**	**-0.27**	**-0.34**	**-0.35**	**-0.12**	**-0.58**
MLS	**0.44**	**0.77**	**0.74**	**0.79**	**0.79**	**0.65**	**0.66**	**0.65**	**0.64**
MEX	**0.05**	**0.14**	**0.27**	**0.14**	**0.15**	**0.25**	**0.15**	**0.16**	**0.26**
PHI	**0.03**	**-0.32**	**0.01**	**-0.32**	**-0.30**	**-0.26**	**-0.41**	**-0.16**	**-0.08**
SAF	**0.25**	**0.61**	**0.29**	**0.62**	**0.58**	**0.74**	**0.79**	**0.80**	**-0.08**

	BEL	**CAN**	**FRA**	**ITA**	**MLT**	**NOR**	**USA**
BEL	1.00						
CAN	0.46	1.00					
FRA	0.74	0.79	1.00				
ITA	0.71	0.28	0.41	1.00			
MLT	0.62	0.73	0.85	0.24	1.00		
NOR	0.54	0.69	0.86	0.34	0.80	1.00	
USA	0.46	0.85	0.77	0.26	0.66	0.62	1.00
CRO	**-0.08**	**-0.28**	**-0.09**	**-0.19**	**-0.13**	**-0.16**	**-0.19**
HKG	**0.48**	**0.85**	**0.76**	**0.29**	**0.70**	**0.64**	**0.96**
HUN	**0.13**	**-0.17**	**0.08**	**0.09**	**-0.06**	**-0.08**	**0.01**
MLS	**0.45**	**0.63**	**0.66**	**0.25**	**0.57**	**0.56**	**0.59**
MEX	**0.20**	**0.58**	**0.49**	**0.12**	**0.48**	**0.34**	**0.60**
PHI	**0.05**	**0.63**	**0.43**	**-0.01**	**0.41**	**0.26**	**0.60**

Correlations for Croatia are based on the subsample $2003M01 - 2009M06$, and correlations for Italy on the subsample $1999M01 - 2009M06$. Boldface entries are correlations between AEs and EMEs risk premia. Missing countries are those for which some interest rate data are missing and, as a consequence, one cannot compute risk premiums for them.

Table B.7: Share of total variation explained by the first three principal components in a Principal Component Analysis of yields across countries and by maturity (using the alternative measure of forward risk premiums)

	3-Month	6-Month
Principal Component 1	49,46	45,71
Principal Component 2	19,48	20,80
Principal Component 3	12,51	14,61

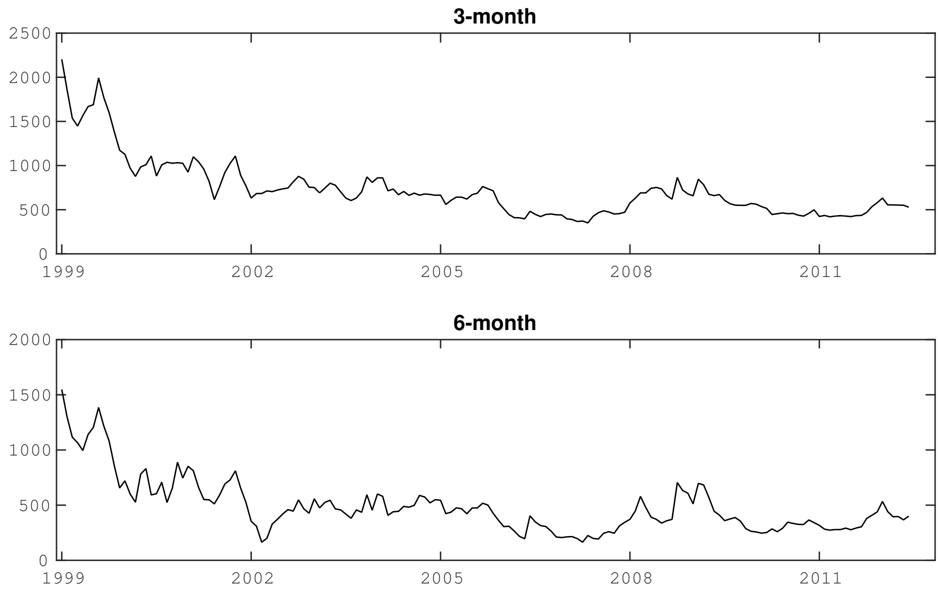

Figure B.1: Time series of maximum differentials by maturity based on the alternative measure of risk premiums (in basis points)

Table B.8: Regression results (Maximum differentials with the alternative measures of forward risk premiums).[c]

	fp(3)		fp(6)			
	1999M01-2002M09	2002M10-2012M06	1999M01-2001M12	2002M01-2005M11	2005M12-2009M10	2009M11-2012M06
Constant	1814.40***	880.80***	1121.96***	-142.48	873.54***	-466.30***
	(293.34)	(88.51)	(360.34)	(120.32)	(214.43)	(172.00)
Trend	**-35.46*** **	**-3.08*** **	**-23.99*** **	**9.33*** **	**-5.11*** **	5.32***
	(5.50)	**(0.92)**	**(4.15)**	**(1.74)**	**(1.79)**	(1.37)
Vix	16.42***	4.02***	16.89**	7.99***	7.57***	0.08
	(3.82)	(1.39)	(6.74)	(1.76)	(1.42)	(2.65)
Fedfunds	-87.44***	-33.74**	-65.07**	-64.36***	-46.84***	147.39
	(29.38)	(13.38)	(27.12)	(18.14)	(11.99)	(499.51)
Obs.	162		162			
R2	0.88		0.88			

Numbers in parentheses are robust standard errors.
*p < 10%. **p < 5%. ***p < 1%.

B.3 Additional Results on the Analysis of the Determinants of Risk Differentials

Table B.9: Determinants of risk premium differentials: 3-month maturity (using the alternative measure of risk premium).

	(1)	(2)	(3)	(4)	(5)
polrisk	24.12	-8.30	-83.23	-231.23	
	(481.57)	(538.62)	(495.99)	(412.29)	
polrisk*dum2002		-214.16			
		(521.14)			
polrisk*dum2008					
polrisk*vix25			26.15		
			(68.57)		
liquid	-2.61*	-0.87	-0.94	-4.13**	-2.91***
	(1.32)	(1.27)	(1.09)	(1.20)	(0.72)
liquid*dum2002		-0.06			
		(1.41)			
liquid*dum2008					
liquid*vix25			-1.35		
			(0.90)		
inflation	42.12**	20.71***	44.54**	42.01**	26.22***
	(15.15)	(4.63)	(16.75)	(14.38)	(4.14)
inflation*dum2002		38.40***			23.54
		(9.23)			(12.41)
inflation*dum2008					
inflation*vix25			-0.62		
			(5.68)		
growth	-5.52	-5.59**	-3.63	-4.30	-0.97
	(5.52)	(2.45)	(6.16)	(6.43)	(3.81)
growth*dum2002		5.80			
		(8.54)			
growth*dum2008					
growth*vix25			-6.97		
			(5.42)		
reer	269.26	-191.75	221.62	239.40	143.23
	(181.24)	(202.07)	(203.84)	(183.39)	(203.76)
reer*dum2002		571.87***			113.93
		(134.31)			(60.72)
reer*dum2008					
reer*vix25			11.06		
			(62.54)		
vix	2.11	-3.25	1.19	-49.92	
	(1.57)	(3.43)	(1.86)	(42.78)	
vix*dum2002		-20.95			
		(12.64)			
vix*dum2008					
fedfunds	-0.94	37.14	-1.22	-228.54**	-19.45
	(12.94)	(19.26)	(13.50)	(84.90)	(11.62)
fedfunds*dum2002		-439.96			
		(401.50)			
fedfunds*dum2008					
euro				43.53	
				(33.65)	
europe				137.40**	153.78***
				(41.46)	(38.56)
nafta				152.17**	142.66**
				(53.10)	(41.46)
emeurope				-49.92	
				(42.78)	
emeasia				-228.54**	-214.14**
				(84.90)	(101.69)
dum2002		-439.96			
		(401.50)			
dum2008					
Constant	-152.37	410.34	-7.38	83.21	49.06
	(431.01)	(446.24)	(442.69)	(391.43)	(225.16)
N	3181	3181	3181	3181	3181
Overall R-Square	0.41	0.51	0.44	0.58	0.60
Wald Chi-Square	35.28***	419.95***	1421.62***	361.36***	1832.29***

The model is estimated over the time period 1999Q1-2010Q1. Model (1) is the baseline specification given by equation (13); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectively. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

Table B.10: Determinants of risk premium differentials: 6-month maturity (using the alternative measure of risk premium).

	(1)	(2)	(3)	(4)	(5)
polrisk	-208.72	-218.01*	-304.33	-438.50	-463.00**
	(293.23)	(89.21)	(284.24)	(226.73)	(142.73)
polrisk*dum2002		56.04			
		(301.18)			
polrisk*dum2008		-385.51			
		(366.18)			
polrisk*vix25			41.03		
			(51.13)		
liquid	-1.70	-0.96	-0.85	-2.70*	
	(0.85)	(1.50)	(0.76)	(1.08)	
liquid*dum2002		1.20			
		(2.24)			
liquid*dum2008		-0.14			
		(1.15)			
liquid*vix25			-0.76		
			(0.60)		
inflation	27.25*	31.04**	30.70*	26.60*	34.35**
	(12.29)	(8.43)	(12.51)	(11.33)	(9.26)
inflation*dum2002		15.18			
		(9.13)			
inflation*dum2008		-21.92*			-28.81**
		(10.20)			(7.76)
inflation*vix25			0.20		
			(5.36)		
growth	-5.84	-10.54**	-5.70	-4.40	-5.24
	(3.64)	(4.06)	(3.68)	(4.54)	(2.84)
growth*dum2002		1.44			
		(6.31)			
growth*dum2008		-9.13			
		(7.08)			
growth*vix25			-7.43		
			(4.65)		
reer	78.52	-267.50*	40.18	33.57	-188.70
	(143.08)	(116.05)	(154.27)	(148.91)	(170.33)
reer*dum2002		460.76**			307.89
		(155.51)			(158.81)
reer*dum2008		217.54			
		(156.44)			
reer*vix25			-7.29		
			(49.09)		
vix	2.17	1.67	1.25	-33.54	
	(1.24)	(2.41)	(1.67)	(35.11)	
vix*dum2002		-37.67***			-0.56
		(6.19)			(3.46)
vix*dum2008		67.49**			-0.69
		(19.74)			(1.69)
fedfunds	6.01	56.72**	5.13	-148.19**	-19.11**
	(10.70)	(16.45)	(11.34)	(56.24)	(5.26)
fedfunds*dum2002		-785.61**			50.03**
		(300.89)			(16.66)
fedfunds*dum2008		-56.04			
		(467.67)			
euro				71.91**	64.26***
				(19.64)	(14.11)
europe					
nafta				61.17**	69.55***
				(15.41)	(12.92)
emeurope				-33.54	-24.52
				(35.11)	(14.40)
emeasia				-148.19**	-125.14**
				(56.24)	(38.83)
dum2002		-785.61**			-468.63*
		(300.89)			(227.50)
dum2008		-56.04			
		(467.67)			
Constant	176.96	711.32**	309.57	451.00	787.76**
	(259.43)	(178.11)	(267.38)	(262.38)	(234.28)
N	2151	2151	2151	2151	2151
Overall R-Square	0.41	0.51	0.44	0.58	0.60
Wald Chi-Square	120.06***	22269.21***	284.77***	381.84***	522.11***

The model is estimated over the time period 1999Q1-2010Q1. Model (1) is the baseline specification given by equation (13); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

Table B.11: Determinants of risk premium differentials: 3-month maturity (Two-Stage Least Square estimators).

	(1)	(2)	(3)	(4)	(5)
polrisk	-117.94	-515.10	-174.50	-370.34**	-171.84**
	(115.94)	(363.34)	(147.77)	(123.31)	(76.50)
polrisk*dum2002		225.53			
		(145.02)			
polrisk*dum2008					
polrisk*vix25			23.99		
			(58.11)		
liquid	-2.22**	16.25	-1.79**	-1.72	-3.30***
	(0.84)	(47.15)	(0.80)	(0.98)	(0.57)
liquid*dum2002		-0.26			
		(3.49)			
liquid*dum2008					
liquid*vix25			0.78		
			(0.43)		
inflation	10.02	-53.18	9.81	11.00	
	(6.24)	(53.83)	(6.86)	(6.30)	
inflation*dum2002		65.58			
		(48.17)			
inflation*dum2008					
inflation*vix25			1.73		
			(4.36)		
growth	-7.15*	-8.01	-6.11	-8.51**	-8.69**
	(3.32)	(7.76)	(5.73)	(3.49)	(3.19)
growth*dum2002		6.65			
		(9.30)			
growth*dum2008					
growth*vix25			-1.87		
			(6.15)		
reer	100.90	-352.39	79.50	90.34	
	(101.38)	(511.26)	(105.74)	(96.65)	
reer*dum2002		525.86			
		(422.73)			
reer*dum2008					
reer*vix25			42.44		
			(39.15)		
vix	-1.03	2.44	-3.61***	55.30	-0.76
	(0.64)	(3.92)	(0.93)	(51.28)	(0.68)
vix*dum2002		10.74			
		(12.99)			
vix*dum2008					
fedfunds	2.21	-15.41	0.24	79.11	
	(4.48)	(15.74)	(4.36)	(41.83)	
fedfunds*dum2002		-855.49			
		(475.99)			
fedfunds*dum2008					
euro				42.75**	31.55**
				(16.01)	(10.81)
europe				-8.59	
				(12.40)	
nafta				-7.04	
				(4.11)	
emeurope				55.30	
				(51.28)	
emeasia				79.11	
				(41.83)	
dum2002		-855.49			
		(475.99)			
dum2008					
Constant	57.45	845.40	168.15	232.73	214.72**
	(215.54)	(752.75)	(237.24)	(206.63)	(84.99)
N	3111	3111	3111	3111	3111
Overall R-Square	0.06	0.01	0.09	0.13	0.05
Wald Chi-Square	547.22***	1843.81***	176.93***	137.32***	107.97***

The model is estimated over the time period 1999Q1-2010Q1. Model (1) is the baseline specification given by equation (13); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

Table B.12: Determinants of risk premium differentials: 6-month maturity (Two-Stage Least Square estimation).

	(1)	(2)	(3)	(4)	(5)
polrisk	-372.77**	-535.35**	-403.86**	-353.38***	-470.48**
	(103.72)	(146.75)	(102.12)	(80.67)	(155.60)
polrisk*dum2002		490.93**			585.02**
		(166.43)			(150.09)
polrisk*dum2008		52.02			
		(124.57)			
polrisk*vix25			89.36		
			(75.53)		
liquid	-1.59*	-1.97	-1.72*	-1.57	-1.85*
	(0.76)	(1.13)	(0.80)	(1.05)	(0.82)
liquid*dum2002		2.51			
		(1.88)			
liquid*dum2008		0.91			
		(0.81)			
liquid*vix25			1.55**		2.36**
			(0.53)		(0.61)
inflation	16.92	5.33	19.02	15.62	
	(8.44)	(14.15)	(10.41)	(9.78)	
inflation*dum2002		17.65			
		(14.30)			
inflation*dum2008		-0.52			
		(12.56)			
inflation*vix25			-9.20		
			(7.49)		
growth	-8.58	7.47	-1.23	-12.23*	-11.04*
	(6.18)	(18.87)	(9.79)	(5.28)	(4.40)
growth*dum2002		-7.96			
		(18.98)			
growth*dum2008		-20.58			
		(18.43)			
growth*vix25			-16.62		
			(10.77)		
reer	109.19	-7.61	99.46	83.65	
	(69.34)	(130.47)	(89.30)	(85.19)	
reer*dum2002		242.45			
		(152.66)			
reer*dum2008		-86.92			
		(277.08)			
reer*vix25			2.43		
			(73.47)		
vix	-1.38	4.93	-3.12*	-6.89	-1.69*
	(1.08)	(3.00)	(1.53)	(44.13)	(0.84)
vix*dum2002		7.85			
		(7.66)			
vix*dum2008		2.00			
		(7.58)			
fedfunds	11.30*	-101.48**	10.38*	43.66	11.67*
	(4.68)	(26.75)	(4.97)	(31.80)	(4.55)
fedfunds*dum2002		-777.63**			4.28
		(260.43)			(7.62)
fedfunds*dum2008		-24.35			
		(346.71)			
euro				18.54	
				(10.24)	
europe					
nafta				-5.49	
				(14.68)	
emeurope				-6.89	
				(44.13)	
emeasia				43.66	
				(31.80)	
dum2002		-777.63**			-486.56**
		(260.43)			(128.16)
dum2008		-24.35			
		(346.71)			
Constant	250.90	549.47	297.60	245.43	478.29**
	(177.32)	(273.39)	(189.09)	(201.68)	(149.70)
N	2114	2114	2114	2114	2114
Overall R-Square	0.12	0.01	0.09	0.13	0.05
Wald Chi-Square	167.39***	175.20***	268.74***	308.17***	81.81***

The model is estimated over the time period 1999Q1-2010Q1. Model (1) is the baseline specification given by equation (13); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.

Table B.13: Determinants of risk premium differentials: 5-year maturity (Two-Stage Least Square estimation).

	(1)	(2)	(3)	(4)	(5)
polrisk	-660.37**	-200.31	-646.27*	-202.23	-332.62**
	(200.61)	(178.57)	(223.50)	(180.26)	(66.85)
polrisk*dum2002		-265.87			
		(188.58)			
polrisk*dum2008		896.52*			943.84*
		(302.88)			(312.28)
polrisk*vix25			99.19		
			(75.90)		
liquid	-6.09**	-0.20	-4.58	-4.64*	-2.03
	(1.83)	(1.63)	(3.42)	(1.83)	(1.15)
liquid*dum2002		-9.57*			-8.74
		(3.74)			(3.91)
liquid*dum2008		-0.36			
		(2.56)			
liquid*vix25			-5.00		
			(2.53)		
inflation	-16.27***	-46.06***	-18.67***	-11.45**	-45.20***
	(1.29)	(2.56)	(2.76)	(2.34)	(2.78)
inflation*dum2002		29.23***			30.63***
		(4.26)			(3.49)
inflation*dum2008		65.98**			61.61***
		(12.47)			(8.84)
inflation*vix25			3.23		
			(2.98)		
growth	-25.64**	34.05**	-16.21*	-34.91**	34.75**
	(4.83)	(10.62)	(5.83)	(9.02)	(8.27)
growth*dum2002		-44.15*			-35.75**
		(15.55)			(10.50)
growth*dum2008		-75.65*			-74.35***
		(24.39)			(12.47)
growth*vix25			-27.42*		-13.84**
			(8.88)		(3.66)
reer	59.73	153.63	113.86*	56.58	127.52
	(29.42)	(85.27)	(36.32)	(43.34)	(63.21)
reer*dum2002		-79.06			
		(91.89)			
reer*dum2008		212.75			
		(239.85)			
reer*vix25			-108.32		
			(58.31)		
vix	0.82	-1.97	3.73	-62.30	
	(3.00)	(1.37)	(4.26)	(42.43)	
vix*dum2002		-24.85			
		(12.29)			
vix*dum2008		46.39			
		(20.00)			
fedfunds	17.47**	83.38**	19.23**	111.20	-30.15*
	(3.59)	(25.92)	(4.46)	(71.20)	(11.33)
fedfunds*dum2002		697.35**			50.40*
		(188.81)			(18.98)
fedfunds*dum2008		-1026.67*			76.07**
		(434.34)			(22.32)
euro				23.60	
				(24.60)	
europe				-8.45	
				(21.62)	
nafta				-67.87***	-56.54**
				(8.17)	(10.70)
emeurope				-62.30	
				(42.43)	
emeasia				111.20	
				(71.20)	
dum2002		697.35**			-9.17
		(188.81)			(71.52)
dum2008		-1026.67*			-874.72*
		(434.34)			-(874.72)
Constant	516.37**	-51.38	380.01*	48.90	179.50
	(119.50)	(137.34)	(136.67)	(108.26)	(118.74)
N	2112	2112	2112	2112	2112
Overall R-Square	0.29	0.60	0.33	0.42	0.59
Wald Chi-Square	3457.76***	3294.54***	373.90***	465.40***	17693.73***

The model is estimated over the time period 1999Q1-2010Q1. Model (1) is the baseline specification given by equation (13); extension (2) includes structural breaks; extension (3) considers the amplifying effects of high volatility; extension (4) includes regional effects; and extension (5) considers all variables that are significant in the previous specifications. Variables *euro, europe, nafta* refer to regional dummies that take the value 1, if an AE belong to the EMU, EU, or NAFTA, respectivelly. Variables *emeurope, emeasia* do the same for EMEs from Europe and Asia respectively. Numbers in parentheses are two-way cluster-robust standard errors. Asterisks indicate significance levels as follows: $*p < 10\%$; $**p < 5\%$; $***p < 1\%$.