Dynamic Responses to Oil Price Shocks: Conditional vs Unconditional (A)symmetry

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Abstract

The impulse-response-function-based Wald test has been gaining wide popularity among researchers seeking to formally test for (a)symmetries in dynamic responses of various macroeconomic aggregates to oil price shocks. However, because the IRF-based Wald test is *conditional* on the magnitude of an oil price shock, it can sometimes prove to be impractical, especially when producing contrasting evidence for shocks of different sizes. To circumvent this problem, this paper suggests considering a nonparametric IRF-*density*-based test in addition to the Wald. The former allows the analysis of (a)symmetries in dynamic impulse responses to positive and negative oil price shocks of a wide range of magnitudes. The test permits inference about a general tendency of (a)symmetries in impulse responses as opposed to (a)symmetries pertinent to a shock of a given size only. The examined (a)symmetry is thus *unconditional* of the magnitude of a shock. Importantly, the testing procedure allows accounting for the relative likelihood of observing the disturbance of a given size.

Keywords: Asymmetry, Impulse Response, Job Creation and Destruction, Macroeconomy, Nonlinearity, Oil Price, SVAR

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1 Introduction and Motivation

Until not long ago, the consensus view among scholars has been that the relationship between the oil price and the U.S. economy is asymmetric and nonlinear. However, this conventional wisdom has recently been challenged, reigniting an interest in studying this relationship. Kilian and Vigfusson (2011a) show that most studies that report asymmetries rely on the use of censored oil price variables in the VAR models which leads to biased parameter estimates and inconsistent estimates of impulse response functions (IRF). More importantly, Kilian and Vigfusson (2011a) make a compelling case against the use of traditional slope-based tests as a formal way to test for (a)symmetry in the structural relationship between the oil price and a macroeconomic aggregate. They argue at length that slope-based tests, by their nature, can provide inference only about (a)symmetries in slope coefficients of the structural equations and not in dynamics responses (to an oil price shock) that are of the primary concern for economists. Rejection/acceptance of symmetry based on slope-based tests does not imply the asymmetric/symmetric IRFs because the latter are highly nonlinear functions of the slope parameters and innovation variances. Slope-based tests are not sufficient (nor are they necessary) for inference about (a)symmetry in impulse responses. Instead, Kilian and Vigfusson (2011a) propose to formally examine the (a)symmetry in responses to oil shocks using the IRF-based Wald test, which has since been gaining wide popularity among researchers working in the field.

Despite its advantages, the IRF-based Wald test falls short at least in one dimension: it is *conditional* on the magnitude of an oil price shock. One can only use this test for the inference about (a)symmetries in impulse responses to a shock of a given magnitude. This can sometimes prove to be impractical, especially in instances when the test provides contrasting evidence for shocks of different sizes. For instance, Kilian and Vigfusson (2011b, p.351) find that, while the IRF-based Wald test[1] provides no support in favor of asymmetric responses of the real U.S. GDP to one-s.d. oil price shocks at the one-year horizon, it however rejects the null of symmetry in the case of two-s.d. shocks. Similarly, Herrera et al. (2015) find that the IRF-based Wald test lends strong support to the asymmetry in one-year responses of the U.S. industrial production to two-s.d. oil price shocks,[2] but they fail to reject the null of symmetry for shocks of a smaller magnitude.

To circumvent the above problem, this paper suggests considering a nonparametric IRF-*density*-based test in addition to the Wald. The former allows the analysis of (a)symmetries in dynamic impulse responses to positive and negative oil price shocks of a wide *range* of magnitudes. Hence, the test permits inference about a general tendency of (a)symmetries in impulse responses, not about (a)symmetries pertinent to a shock of a given size only. Thus, the examined (a)symmetry in IRFs is *unconditional* of the magnitude of a shock, which is a naturally desired property. Building upon the work of Li et al. (2009), the idea is to compare empirical densities of impulse responses to positive and negative oil price shocks of a range of magnitudes. Importantly, the testing procedure allows accounting for the relative likelihood of observing the disturbance of a given size.

I showcase the practical convenience of the IRF-*density*-based test for studying (a)symmetries in dynamic responses of macroeconomic aggregates to oil price shocks using the example of job flows in U.S. manufacturing sector in the 1973–2005 period. Despite that costly labor reallocation has been one of the most widely cited propagation mechanism contributing to asymmetry in responses to oil price shocks, very little applied work has been done on effects of oil price shocks on labor market flows, except for studies by Davis and Haltiwanger (2001), Herrera and Karaki (2012) and Malikov (2012). Using the IRF-based Wald tests, Herrera and Karaki (2012) report no evidence of asymmetry in responses of job flows to one-s.d. oil price shocks, whereas some evidence is found for two-s.d. innovations. Given this contrasting evidence, I use the IRF-density-based test to take another close

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[1] Using Hamilton's (2003) three-year net oil price increase measure to model potential nonlinearities.

[2] When using Hamilton's (1996) one-year net oil price increase measure.

look at manufacturing job flows. I find that, when accounting for the historical distribution of oil price innovations, the U.S. manufacturing job flows appear to respond to oil price shocks asymmetrically.

2 Econometric Strategy

This paper uses a trivariate *block*-structural VAR system, along the lines of that in Kilian and Vigfusson (2011a), which allows for potential asymmetries and nonlinearities in the relationship between the real oil price and gross job flows:

$$x_t = \beta_{10} + \sum_{j=1}^{p} \beta_{11,j} x_{t-j} + \sum_{j=1}^{p} \beta_{12,j} y_{1,t-j} + \sum_{j=1}^{p} \beta_{13,j} y_{2,t-j} + \epsilon_{1t} \tag{1a}$$

$$y_{1,t} = \beta_{20} + \sum_{j=0}^{p} \beta_{21,j} x_{t-j} + \sum_{j=1}^{p} \beta_{22,j} y_{1,t-j} + \sum_{j=1}^{p} \beta_{23,j} y_{2,t-j} + \sum_{j=0}^{p} \gamma_{21,j} x^*_{t-j} + \epsilon_{2t} \tag{1b}$$

$$y_{2,t} = \beta_{30} + \sum_{j=0}^{p} \beta_{31,j} x_{t-j} + \sum_{j=1}^{p} \beta_{32,j} y_{1,t-j} + \sum_{j=1}^{p} \beta_{33,j} y_{2,t-j} + \sum_{j=0}^{p} \gamma_{31,j} x^*_{t-j} + \epsilon_{3t}, \tag{1c}$$

where x_t is the log difference of the real oil price; x^*_t is some nonlinear transformation of the real oil price; y_{1t} and y_{2t} are the growth rates of gross job creation and destruction, respectively; ϵ_{1t}, ϵ_{2t} and ϵ_{3t} are zero-mean serially uncorrelated disturbances. Gross job creation is defined as the net sum of employment gains at expanding and entering establishments, while gross job destruction is said to be the net sum of employment losses at contracting and exiting establishments. Naturally, the difference between gross job creation and destruction is the net job change (growth).

I consider two popular nonlinear transformations of the oil price that censor price declines as well as minor price increases: the net oil price increase over the prior four-quarter maximum (Hamilton, 1996) and twelve-quarter maximum (Hamilton, 2003), respectively defined as

$$x^4_t = \max\{0, o_t - \max[o_{t-1}, \ldots, o_{t-4}]\} \tag{2}$$

$$x^{12}_t = \max\{0, o_t - \max[o_{t-1}, \ldots, o_{t-12}]\}, \tag{3}$$

where o_t is the log of the real oil price.[3]

Model (1) consists of two (structural) blocks: *(i)* the real oil price and *(ii)* gross job flows. The equations imply a cross-block identification of two sorts. First, a block-recursive ordering is applied to the VAR system with the oil price (the first block) assumed to be contemporaneously exogenous to real economic activity, namely gross job flows (the second block). This is a widely accepted assumption in the literature which has also been validated empirically (Kilian and Vega, 2011). Such ordering localizes a structural innovation in the real oil price in ϵ_{1t}, rendering the latter as an oil price shock. Second, equations (1b)–(1c) allow for potential nonlinearities in the relationship between the oil price and gross job flows by including the nonlinear transformation of the price of oil. Note that, unlike Herrera and Karaki (2012), I do not seek identification within the second block of the system. In the dimension of gross job flows, model (1) is estimated in its reduced form, rendering ϵ_{2t} and ϵ_{3t} as (partially) reduced-form innovations in gross job creation and destruction. This approach is similar to that of Davis and Haltiwanger (2001) and is justified on several grounds. First, identification of the structural disturbances to the real oil price (ϵ_{1t}) alone is sufficient for the investigation of

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[3] The oil price increase, which does not censor minor price increases, is another popular choice for x^*_t. To conserve space, the results for this nonlinear transformation are relegated to Supplementary Appendix D.

potential asymmetries in responses of job flows to oil price shocks. Second, the choice of a proper identification scheme for gross job creation and destruction is not trivial. Imposing an exclusion restriction on either of the two variables, which Herrera and Karaki (2012) resort to, may be too restrictive and counterintuitive provided that both processes are simultaneous in nature. While the sign, parameter range and/or long-run neutrality restrictions à la Davis and Haltiwanger (1999) may be a more viable identification strategy, it however would play no role in studying responses to already-identified oil price shocks which is the main focus of this paper.[4]

2.1 Estimation Details and Data

The sample period examined in this paper is 1973:I–2005:I. Kilian and Vigfusson (2011b) argue at length that the post-1973 period is more appropriate for the analysis of the structural dynamic relationship between the real oil price and economy. The choice of the endpoint is rather dictated by the availability of the job flow data. Consistent with Hamilton (2003), Kilian and Vigfusson (2011b), Herrera and Karaki (2012) and Malikov (2012), the lag order of a VAR system is set to four ($p = 4$).

For the nominal price of oil, this paper uses Refiners' Acquisition Cost (RAC) for imported crude oil. Since the RAC series is not available for the period prior to 1974:I, I follow Kilian (2009) and extrapolate the time series back to 1973:I at the growth rate of PPI. To obtain the real price of oil, I deflate the above series by the composite CPI index. I opt for the real price of oil (over nominal) not only because doing so is consistent with economic theory but also because the use of nominal oil price in the post-1973 period may be ill-advised, since the latter likely reflects inflation innovations (Kilian and Vigfusson, 2011b).

The series of gross job creation and destruction for the U.S. manufacturing sector come from Davis et al. (2006). These are seasonally adjusted quarterly growth rates that are the employment-weighted means of the plant-level job growth rates calculated using the average of the current and prior quarters' employment at the plant in denominator. This measure of job flows is symmetric about zero and bounded from both sides with exit and entrance of a plant corresponding to boundaries of the [–2, 2] interval.

3 *Conditional* Asymmetry in Impulse Responses

Given the nonlinearity of model (1), the dynamic impulse responses of job flows to an oil price innovation are *not* invariant to the magnitude of the shock. Consequently, the degree of asymmetry in IRFs is likely to be highly dependent (conditional) on the magnitude of innovations.

Based on model (1), I estimate impulse responses of gross job creation and destruction to a positive and negative oil price shock of a given magnitude at an eight-quarter horizon. I also construct impulse responses of net job growth as the difference between the estimated responses of job creation and destruction. Following Kilian and Vigfusson (2011a), I start by disturbing the model with shocks of two sizes set equal to one and two standard deviations of the oil price residuals. Naturally, in order to examine potential asymmetries in impulse responses I subject the system to a positive and negative oil shock. One can formally test for asymmetries in the obtained impulse responses using the IRF-based Wald test proposed by Kilian and Vigfusson (2011a, 2011b). Specifically, I test the

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[4] Model (1) does not allow for an endogenous monetary policy response to oil price shocks as in Davis and Haltiwanger (2001). Here, I opt for a more parsimonious model, given the growing evidence casting doubt on the importance of monetary policy responses to oil price changes. Even if model (1) is modified to incorporate monetary policy (proxied by the federal funds rate), I find that, regardless of the ordering assumption, the results are qualitatively unchanged.

Table 1. IRF-based Wald Tests of (Conditional) Asymmetry

	$x_t^* = x_t^4$			$x_t^* = x_t^{12}$		
Horizon	JC	JD	NG	JC	JD	NG
			One-s.d. Shock			
0	0.073	0.051	0.051	0.597	0.493	0.520
1	0.096	0.103	0.097	0.572	0.785	0.725
2	**0.034**	0.170	0.052	0.757	0.915	0.886
3	**0.022**	0.092	0.077	0.881	0.972	0.830
4	**0.042**	0.077	0.134	0.917	0.965	0.914
5	0.053	0.066	0.184	0.960	0.985	0.951
6	0.081	**0.009**	**0.008**	0.935	0.993	0.932
7	0.083	**0.009**	**0.004**	0.959	0.991	0.952
8	**0.034**	**0.011**	**0.005**	0.978	0.990	0.927
			Two-s.d. Shock			
0	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
1	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
2	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
3	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
4	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
5	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
6	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
7	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**
8	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**

The reported are *p*-values with the 5% significance in bold.

null hypothesis of *jointly* symmetric impulse responses:

$$\mathbb{H}_0 : I_y\left(h, \delta\right) = -I_y\left(h, -\delta\right) \qquad \forall\ h = 0, \ldots, H \tag{4}$$

for a *given* size of an oil price shock δ, where $I_y\left(h, \delta\right)$ is the response of y_t (job creation, destruction or net job growth) to an innovation in x_t of magnitude δ at horizon h.[5] The asymptotically χ^2 distributed IRF-based Wald test statistic of joint symmetry is given by

$$\mathcal{W} = \left(\mathbf{R}\widehat{\mathbf{I}}\right)' \left(\mathbf{R}\widehat{\mathbf{\Omega}}\mathbf{R}'\right)^{-1} \left(\mathbf{R}\widehat{\mathbf{I}}\right) \sim \chi^2_{H+1}, \tag{5}$$

where $\widehat{\mathbf{I}} = \left[\widehat{I}_y(0, \delta), \ldots, \widehat{I}_y(H, \delta), \widehat{I}_y(0, -\delta), \ldots, \widehat{I}_y(H, -\delta)\right]'$ is the $2(H+1) \times 1$ vector of impulse response estimates, $\widehat{\mathbf{\Omega}}$ is the $2(H+1) \times 2(H+1)$ recursive-design bootstrap estimate[6] of the covariance matrix $\mathbb{E}[(\widehat{\mathbf{I}} - \mathbf{I})(\widehat{\mathbf{I}} - \mathbf{I})']$, and $\mathbf{R}$ is the $(H+1) \times 2(H+1)$ "design" matrix defined as $\mathbf{R} \equiv [\mathbf{I}_{H+1}\ \mathbf{I}_{H+1}]$ with $\mathbf{I}_{H+1}$ being the identity matrix of dimension $(H+1)$. For details on the computation of the IRFs and test statistics, see Supplementary Appendices A and B.

Table 1 reports p-values for the above IRF-based Wald test performed on impulse responses of gross job creation and destruction, and net job growth.[7] Using either of the two nonlinear oil price transformations, I find strong evidence of asymmetric responses of manufacturing job flows to a large (two-s.d.) oil price shock only. Considering shocks of a smaller magnitude (one s.d.) however significantly weakens the case for asymmetry. Particularly, when modeling nonlinearities by means of x_t^{12}, I fail to reject the null of symmetry for all job flows at *all* horizons. In the case of x_t^4, the data similarly provide little strength against the null of joint symmetry, uncovering a rather weak evidence of significant asymmetries at longer horizons only.

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[5] $I_y\left(h, \delta\right)$ is the unconditional impulse response obtained via integrating over the set of possible prior histories.

[6] Computed over 149 iterations.

[7] All the estimated IRFs are jointly significantly different (at the 5% level) from zero at a two-year horizon.

4 *Unconditional* Asymmetry in Impulse Responses

From Section 3, we see that the examination of asymmetries in impulse responses to one- and two-s.d. oil price shocks using the IRF-based Wald test produces rather contrasting findings. Owing to its dependence on the magnitude of a shock, the Wald test thus falls short. What if one claims that oil shocks of different magnitudes, rather than one and two standard deviations, are more appropriate for the study? This raises a question of how one should determine the magnitude of a shock in responses to which asymmetries are to be studied. How many different shocks should one consider to satisfactorily answer the research question, and what if different shocks point to contrasting conclusions? Indeed, this paper finds evidence of asymmetry in the case of two-s.d. oil price shocks only. Given that the historical likelihood of observing the disturbance of this size or larger is a mere 6%, it is not fully clear if one should interpret the results in favor of a mostly *symmetric* relationship between job flows and the oil price, as suggested by the tests on IRFs to a one-s.d. shock.

I therefore suggest considering another IRF-based test in addition to the Wald. This test allows the investigation of (a)symmetries in job flows' responses to positive and negative oil price shocks of a wide *range* of magnitudes. Thus, the test permits inference about a general tendency of (a)symmetries in impulse responses, not about (a)symmetries pertinent to a shock of a given size only.

Based on model (1), I estimate IRFs of employment flows to positive and negative oil price shocks of many different sizes. Specifically, I consider oil price disturbances of three thousand distinct (absolute) magnitudes that range from a thousandth of a standard deviation to three standard deviations[8] of the oil price residuals from equation (1a). That is, I disturb the structural model (1) with positive and negative oil shocks of the (absolute) magnitudes $\delta \in \left\{\frac{1}{3000}3\sigma, \frac{2}{3000}3\sigma, \dots, \frac{2999}{3000}3\sigma, 3\sigma\right\}$. This produces two empirical distributions of impulse responses to positive and negative oil price disturbances for each job flow at any given horizon.

Under the null of *joint* symmetry in responses of a given employment flow y_t to positive and negative oil price shocks, the joint empirical distribution of impulse responses to positive oil price disturbances $f_1\left[I_y^+(0), \dots, I_y^+(h)\right]$ is equal to the joint empirical distribution of the negatives of responses to negative oil price disturbances $f_2\left[-I_y^-(0), \dots, -I_y^-(h)\right]$ over a given horizon h:

$$\mathbb{H}_0 : f_1\left[I_y^+(0), \dots, I_y^+(h)\right] = f_2\left[-I_y^-(0), \dots, -I_y^-(h)\right] \qquad \forall\ h = 0, \dots, H. \tag{6}$$

The hypothesis about the equality of empirical probability densities directly follows from the definition of a joint symmetry in impulse responses, provided that the system is disturbed by positive and negative shocks of the very same magnitudes appearing with the same frequency,[9] and that the IRFs are monotonic functions of oil shock. To test this hypothesis, I use a recent extension of Li's (1996) univariate density equality test to a multivariate joint density case by Li et al. (2009). The test is based on the integrated square difference, a measure of global closeness between two densities:

$$\mathcal{I} = \int \Big(f_1(z) - f_2(z)\Big)^2 dz = \int \Big(f_1(z)dF_1(z) + f_2(z)dF_2(z) - f_1(z)dF_2(z) - f_2(z)dF_1(z)\Big), \tag{7}$$

where $F_1(\cdot)$ and $F_2(\cdot)$ are the cumulative distribution functions of $\left(I_y^+(0), \dots, I_y^+(h)\right)$ and $\left(-I_y^-(0), \dots, -I_y^-(h)\right)$, respectively. Li et al. (2009) propose the following asymptotically normally distributed test statistic:

$$\mathcal{T}_n = \frac{N\left(\lambda_0 \cdots \lambda_H\right)^{1/2}}{\widehat{\sigma}_n} \widehat{\mathcal{I}}_n \ \sim\ \mathbb{N}(0,1), \tag{8}$$

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[8] I choose three standard deviations as an endpoint since it is about the size of the largest shock in the sample.

[9] Here, the system is disturbed by a positive and negative shock of a given magnitude only once.

Table 2. IRF-Density-based Tests of (Unconditional) Asymmetry

	Uniform Distr. of Shocks			*Emp. Distr. of Shocks*		
Horizon	JC	JD	NG	JC	JD	NG
		$x_t^* = x_t^4$				
0	250.70	99.07	222.94	83.00	77.54	83.28
1	204.13	168.85	185.66	65.62	65.83	71.05
2	154.53	112.73	148.30	73.32	83.59	66.31
3	106.04	85.48	114.91	66.18	60.35	59.13
4	78.92	66.51	79.51	63.73	83.10	59.37
5	63.61	53.18	56.43	61.43	62.31	61.29
6	50.39	42.82	44.84	83.58	59.21	83.01
7	40.55	34.92	36.67	67.44	54.37	60.22
8	35.17	29.23	30.98	68.62	62.24	59.12
		$x_t^* = x_t^{12}$				
0	280.52	284.58	270.31	77.95	74.95	72.17
1	244.94	206.13	216.80	62.11	77.79	71.63
2	168.48	133.74	144.99	67.50	74.01	66.45
3	112.23	90.53	111.15	63.98	83.55	66.83
4	84.68	69.93	84.04	80.51	78.85	71.98
5	71.99	54.14	62.22	66.57	65.03	59.45
6	57.16	45.12	49.59	80.69	59.29	59.13
7	48.04	38.89	40.60	59.12	59.11	83.58
8	41.09	33.03	35.73	66.44	68.11	59.68

The reported are $\mathcal{T}_n$ test statistics.

where N is the number of impulse response estimates;[10] λ_h is the optimal bandwidth for the impulse response at horizon $h = 0, \ldots, H$; $\widehat{\mathcal{I}}_n$ is the kernel estimate of the integrated square difference $\mathcal{I}$; $\widehat{\sigma}_n^2$ is the kernel estimate of the asymptotic variance of $N(\lambda_0 \cdots \lambda_H)^{1/2}\mathcal{I}$. I use second-order Gaussian kernel throughout. The optimal bandwidths are selected via least squares cross-validation. Since the limiting distribution tends to poorly approximate the finite sample distribution of $\mathcal{T}_n$, I use bootstrap critical values obtained from 199 iterations (for details, see Supplementary Appendix C).

The left panel of Table 2 reports the estimates of the test statistic $\mathcal{T}_n$. The test statistics are all large in magnitude with corresponding bootstrap p-values that are consistently less than 10^{-6} (omitted), leading to an overwhelming rejection of the respective null hypotheses of symmetry in impulse responses as far as at a two-year horizon regardless of the nonlinear oil price transformation.

The expected criticism of the above tests is that they do not take into account the non-uniform likelihood of observing oil price shocks of different magnitudes. Recall that in order to generate a family of IRFs, system (1) is disturbed by oil shocks of a given magnitude only once, thus assuming uniformly distributed (magnitudes of) the oil price innovation. However, the historical distribution of oil price shocks is non-uniform: large shocks are less likely to occur than shocks of smaller magnitudes. A straightforward way to account for this would be to consider a distribution of IRFs generated via disturbing the system by oil price shocks of historical magnitudes obtained from the residuals of equation (1a). Such an approach is however unlikely to be reliable due to an acute "curve of dimensionality" problem.[11] I therefore pursue an alternative approach. Specifically, using the inverse transformation method, I construct a (pseudo-)random sample of oil shocks ranging from $\frac{1}{3000}3\sigma$ to 3σ in their absolute magnitude, which has the same distribution as the historical oil price disturbances estimated by equation (1a). I essentially "extrapolate" oil price shocks in order to increase the size of the sample while maintaining the distribution of the oil price innovations.[12] Using the constructed

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[10] The interval from $\frac{1}{3000}3\sigma$ to 3σ gives 3,000 estimates of IRFs to positive and negative oil price shocks each.

[11] Recall that the sample size is 1973:I-2005:IV, yielding 125 (129 quarters minus 4 lags) residuals. These are hardly enough data in order to obtain reliable nonparametric kernel estimates of the H-variate joint densities of the IRFs.

[12] For details, see Supplementary Appendix C.

sample of oil price shocks (absolute magnitudes of thereof, to be precise), I obtain two empirical distributions of impulse responses to positive and negative oil price disturbances for each job flow at any given horizon. Li et al.'s (2009) test is then applied to these distributions. The estimates of the test statistics are reported in the right panel of Table 2. While the numerical values of test statistics have generally decreased compared to those presented in the left panel of the table, all the corresponding bootstrap values are still virtually zero (not reported), thus lending a strong support in favor of *asymmetry* in dynamic responses of U.S. manufacturing job flows to oil price innovations under both nonlinear oil price transformations in general, *regardless* of the magnitude of an oil shock.

5 Conclusion

This paper argues in favor of a nonparametric IRF-*density*-based test that allows testing for (a)symmetries in dynamic impulse responses of macroeconomic aggregates to positive and negative oil price shocks of a wide range of magnitudes. The test permits inference about a general tendency of (a)symmetries in impulse responses as opposed to (a)symmetries pertinent to a shock of a given size only. The examined (a)symmetry is thus *unconditional* of the magnitude of a shock. Importantly, the testing procedure allows accounting for the relative likelihood of observing the disturbance of a given size.

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